Institutional Investment and International Risk-sharing

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ABSTRACT

We develop and estimate a new asset pricing model to study how global institutional investment affects global and local risk premia in 38 markets. By investing across countries, institutional investors facilitate international risk-sharing between home-biased retail investors. This risk-sharing channel depends on the scope of institutional investors' mandate, their risk bearing capacity and substitutability between securities from different countries. Securities earn a global market risk premium as well as an *institutional local risk premium*. In addition, securities that are not invested by institutions earn a *retail local risk premium*. The average annual institutional local risk premium is 2.76% in developed markets and 6.27% in emerging markets. The average annual retail local risk premium is 1.71% in developed markets and 2.65% in emerging markets. Higher firm-level global institutional ownership reduces the cost of capital in emerging markets.

JEL classification: G12, G15, F30, G20.

Keywords: international asset pricing, institutional investors, market integration.

I. Introduction

Institutional investors are increasingly diversifying their portfolios internationally (Faias and Ferreira, 2017). The global institutional ownership of common equity has increased from 2% to 19% between 2000 and 2020.[1] Global institutional investment is an important pass-through for international risk-sharing because retail investors have maintained persistent home bias despite financial globalization.[2] How does global institutional investment affect international risk-sharing and the resulting global and local risk premia? Answers to this question have implications for understanding the effect of global institutional investment on firms' cost of capital. In order to answer this question, we need to extend existing theories of international asset pricing (see, for example, Errunza and Losq (1985) and De Jong and De Roon (2005)) by introducing the different investment scopes of institutional investors and retail investors.

In this paper, we develop and estimate a novel international asset pricing model to study how institutional investment affects global and local risk premia when retail investors are home-biased. Our theory provides a unique decomposition of market-level local risk premium into an institutional component and a retail component. The theoretical framework can be used to understand who drives time-varying risk premia and how does global institutional investment affect firms' cost of capital. We estimate the model and find that both components are economically important across a wide range of countries. In addition, in emerging markets, firms with higher global institutional ownership have lower cost of capital.

Our model features a domestic and a foreign country. Institutional investors invest globally but are constrained by their mandate to invest in a limited set of securities in each country.

[1] We follow Bartram et al. (2015) and define global institutions as institutional investors whose maximum weight in one country does not exceed 90% and whose maximum weight in one region does not exceed 80%. For example, The Vanguard Group, Inc is a global institutional investor whereas Berkshire Hathaway (Investment management) is a non-global institutional investor that is focused in the US market.

[2] On average, households invest more than 80% of their portfolios at home (see Appendix A). Karlsson and Nordén (2007); Seasholes and Zhu (2010) provide direct evidence of the strong local bias of retail investors due to implicit barriers related to cultural and informational environments.

In contrast, retail investors in both countries invest only at home but in all local securities. Markets are partially segmented across country borders as well as between institutional and retail securities in each country. We define securities that are included in the institutional mandate as *institutional securities* and the remaining securities that are only invested by retail investors as *retail securities*.

To understand the risk-sharing mechanism in our model, consider two companies in each country. The Chinese tech company Tencent is an institutional stock with global institutional ownership of 18%. The Chinese fishery company Zoneco is a retail stock with zero global institutional ownership. Global institutions like Vanguard trade institutional securities from different countries and channel risk-sharing across the border between home-biased retail investors. Due to its limited mandate, Vanguard cannot directly invest in the fishery company. However, it can indirectly get partial exposure to the Fishery company through a replicating portfolio consisting of the tech company. To hedge its investment in the tech company, Vanguard would sell a portfolio of US institutional securities like Apple, which optimally mimics the Chinese tech company, to home-biased US retail investors. This mimicking portfolio is the "homemade" substitute portfolio for Chinese investment to home-biased US retail investors. We define the component of investment opportunities in a country that can be replicated by its institutional securities as their *attainable returns*.

The model predicts three sources of risk premia in the equilibrium. All securities earn an *attainable world market risk premium* and an *institutional local risk premium*. The attainable world market risk represents the global risk that is shared through institutional investment. The *institutional local risk premium* arises due to imperfect cross-border risk-sharing. Cross-border risk-sharing is channeled by institutional investment because retail investors from different countries do not trade with each other due to their home bias. This risk-sharing channel depends on the degree of substitutability between institutional securities from different countries.[3] Because foreign institutional securities do not perfectly substitute domestic institutional securities, institutional investors have to bear the residual local risk in domestic institutional securities that cannot be hedged with foreign institutional securities. This residual local risk gives rise to the *institutional local risk premium* that increases with the risk aversion of institutional investors. Retail securities earn an additional *retail local risk premium* because retail investors in each country have to hold the residual component of retail securities that cannot be replicated by local institutional securities. The retail local premium increases with the risk aversion of retail investors. Our theoretical results show that the institutional local premium can be reduced if global institutional investors have higher risk-bearing capacity, and the retail local premium can be reduced if global institutional investors enlarge their investment mandate.

We validate these predictions by estimating the model using monthly individual stock returns data in the Compustat Global database from January 2000 to December 2020 in 38 countries. We first perform a simple Fama and MacBeth (1973) two-pass regression to test whether the institutional local factor and the retail local factor are significantly priced in the cross-section of individual stocks. We find wide-spread evidence in support of the significance of both local factors. The institutional local factor is significantly and positively priced in 15 out of 23 developed markets (DMs) and in 9 out of 15 emerging markets (EMs). The retail local factor is significantly and positively priced in 7 out of 23 DMs and in 6 out of 15 EMs.

To further analyze how global and local risk premia vary over time and across individual stocks, we perform a second estimation using the conditional two-pass regression framework developed by Gagliardini, Ossola, and Scaillet (2016) (GOS). The GOS framework uses common and stock-specific instruments to model time-varying risk premia and factor exposure, which yields estimates for country-level as well as individual stock level time-varying risk pre-

[3] Substitutability means the extent to which a security can be replicated by other securities. In the extreme case of perfect substitutability, there exists a portfolio of foreign securities that perfectly replicates a domestic security. Greenwood, Hanson, and Liao (2018) also use the notion of substitutibility.

mia. We then quantify the effect of global institutional ownership on firms' cost of capital by running panel regressions of model-implied risk premia on lagged firm-level global institutional ownership.

Our analysis yields three main findings: First, we find that both the institutional local and the retail local premia are economically important in both DMs and EMs. The institutional local premium is on average 2.76% in DMs and 6.27% in EMs. The retail local premium is on average 1.71% in DMs and 2.65% in EMs. We use country-level institutional ownership as a proxy for institutional investors' risk-bearing capacity in each country. Consistent with our theory, the institutional local risk premium is lower in countries with higher global institutional ownership. Second, in DMs, time-variation in total risk premium is driven by the attainable world premium and the institutional local premium, whereas in EMs, the retail local risk premium also contributes to variation in total risk premium. Third, after controlling for firm-level variables and country-time fixed effect, we find that a 1% increase in global institutional ownership predicts a reduction in total risk premium of 8.1 bps in EMs. Therefore, global institutional ownership reduces the cost of capital in EMs.

Our work is part of the literature on international asset pricing. A long-standing question in this literature is whether assets are priced globally or locally. The existing literature has motivated local risk premia in two ways. First, some studies motivate market-level local risk premium based on the assumption of a pricing kernel that is linear in the local market return (see, for example Bekaert (1995) and Bekaert et al. (2007)).[4] Second, formal asset pricing theories such as Errunza and Losq (1985), Eun and Janakiramanan (1986) and De Jong and De Roon (2005) show that uninvestable securities earn a local risk premium due to non-sharable orthogonal local risk. These theories have focused on investability at the security level, assuming investors diversify globally into investable securities that are not subject to explicit investment barriers. This assumption leads to the prediction that globally

[4] Bekaert et al. (2007) also show that a local liquidity premium could arise in integrated markets due to Jensen's inequality terms of this pricing kernel formulation.

traded securities should be priced only by global risk factors. Contrary to this prediction, the body of international finance research shows that prices of globally traded assets depend upon country-specific local risk factors (Lewis, 2011).[5] Therefore investability alone does not fully characterize global equity markets structure, and we contribute by considering the actual investment scopes of different types of investors. Our institutional and retail local risk premia are similar to the orthogonal local risk premium in these models, which arises from residual local risk within a set of securities that cannot be spanned by other securities. Instead of constructing local risk factors based on security-level investment restrictions, our institutional and retail local risk premia result from the home bias of retail investors and the mandate constraint of global institutional investors, which allows us to explain why investable securities also earn local risk premium and to analyze the contribution of different types of investors to total local risk premium.

Our framework borrows from the literature on arbitrage activities and market integration where arbitrageurs trade across two segmented markets. Arbitrage activities may be constrained because of frictions such as collateral constraints (Gromb and Vayanos, 2002), holding costs (Tuckman and Vila, 1992) or slow-moving capital (Greenwood, Hanson, and Liao, 2018). Institutional investors act as "arbitrageurs" in our model.[6] We consider a case in which the arbitrageur has a limited mandate. Although retail securities from two countries are only invested by their local retail investors and hence do not share a common investor, institutional investors could indirectly integrate these markets through their correlation with their local institutional securities.

Our empirical analyses contribute to a growing literature that investigates how institutional investment affects asset returns. In a domestic setting, Edelen, Ince, and Kadlec (2016)

[5]There is abundant recent evidence showing that local risk factors are still necessary to explain the cross-section of security returns absent explicit barriers. See Hou, Karolyi, and Kho (2011), Fama and French (2012), Petzev, Schrimpf, and Wagner (2016), Hollstein (2020), Chaieb, Langlois, and Scaillet (2021).

[6]Strictly speaking, unless domestic and foreign institutional securities could perfectly replicate each other, institutional investors could not have a risk-free arbitrage in our model.

document that institutional demand is negatively associated with stock returns in the long run, which cannot be explained by price reversals. Pavlova and Sikorskaya (2020) show that inelastic demand from benchmarked institutional investors predicts lower future stock returns. Unlike these studies that focus on how institutional demand affects domestic asset returns, our focus is on how institutional investment acts as an international risk-sharing channel across segmented markets and how it affects both global and local risk premia. Studies about institutional investment and international asset pricing are sparse. Bartram et al. (2015) show that the pricing of stocks can be explained by their co-movement with foreign stocks sharing similar institutional investors. Faias and Ferreira (2017) document that industry and global factors are more important than country factors in explaining the return variations of stocks with higher institutional ownership. While their empirical analyses speak to how institutional ownership affects the proportion of variance explained by global versus local factors, our focus is instead on quantifying how institutional ownership affects the levels of global and local risk premia. Kacperczyk, Sundaresan, and Wang (2021) show that foreign institutional ownership reduces firms' cost of equity by increasing stock price informativeness. Our paper, on the other hand, show that global institutional ownership reduces cost of capital by improving cross-border risk-sharing.

Section II develops an asset pricing model with mandate-constrained global institutional investors and home-biased retail investors; Section III discusses the empirical framework; Section IV presents the empirical results and V concludes.

II. Asset Pricing with Global Institutions and Local Retail Investors

A. Model setup and assumptions

We consider an economy with two countries, domestic (D) and foreign (F). In this section and in Appendix B, we use bold font to represent a vector, $'$ to denote the transposition operator, superscript * to represent foreign securities.

A1: (No currency risk) We follow the literature of international asset pricing with barriers (see, e.g. Stulz (1981)) and assume that the purchasing power parity (PPP) holds and that there is no currency risk. All security returns are denominated in the domestic currency.

A2: (Agents) There are three types of investors: *institutional investors* are global investors who invest in both countries, *domestic retail investors* and *foreign retail investors* have very strong home bias and invest only in their home country.[7] In addition, we assume that institutional investors are constrained by their investment mandate to consider only a subset of securities from each country. This subset of securities is the choice set of institutional investors defined in Koijen, Richmond, and Yogo (2022) and we assume that it is exogenous.[8] We also use *choice set* to refer to the set of securities that an investor considers when making her portfolio choice.[9] The choice sets of retail investors consist only of their home securities. Without loss of generality, we work with one representative institutional investor i, one representative domestic retail investor d and one representative foreign retail investor f.

A3: (Single-period portfolio choice). Agents have CARA preference. Investors receive an endowment at time t, trade at time t to maximize utility over their wealth at time

[7]For simplicity, we abstract away potential differences between domestic and foreign institutional investors and view them as a common type. The institutional investor in the model can be viewed as large global financial institutions that trade across country borders. Local institutions who only invest locally could not be distinguished from retail investors. Focusing on large global institutions is a reasonable assumption because the domestic institutional sector is relatively small in most of our sample countries. Domestic institutional ownership is on average 5.99% in DMs and 3.79% in EMs by the end of 2020 (Chaieb, Errunza, and Lu, 2021) and their impact on international risk-sharing is expected to be smaller compared to large global institutions. Chan, Covrig, and Ng (2005) and Lau, Ng, and Zhang (2010) document the home bias of domestic institutional investors around the world.

[8]Koijen, Richmond, and Yogo (2022) make a similar assumption about exogenous choice sets. This assumption is supported by limited inclusion by major global indices. FTSE all world indices, used by global institutions as a major benchmark, include less than 20% of all the stocks in most countries (see Appendix A). Due to optimized sampling, actual number of firms in the benchmark invested by global institutions can be even smaller (Pavlova and Sikorskaya, 2020). The limited investment scope of global institutional investors can be motivated by limited information (Merton, 1987), regulatory constraints, environmental, social and governance (ESG) concerns (Matos, 2020) and information costs (De Marco, Macchiavelli, Valchev, et al., 2018).

[9]The choice set can be thought of as the preferred habitat of investors due to implicit barriers that arise from many reasons including cultural and familiarity bias, see Hollstein (2020).

$t+1$. Investor $k \in \{d, f, i\}$ has absolute risk aversion γ^k and solves the following single period portfolio choice problem from time t to $t+1$:[10]

$$\max_{x^k_{j\in\mathcal{C}^k}} \mathbb{E}[U(W^k_{t+1})] = \max_{x^k_{j\in\mathcal{C}^k}} \mathbb{E}[-\exp(-\gamma^k W^k_{t+1})]$$
$$W^k_{t+1} = W^k_t(1+r_f) + \sum_{j\in\mathcal{C}^k} x^k_j r_j$$

where investor k chooses her optimal dollar investment x^k in securities in her choice set $\mathcal{C}^k$.

A4: (Securities and market structure) There are N risky securities and one risk-free security. A market segment is a set of securities that share the same group of investors. Home bias of retail investors and the limited mandate of the institutional investor give rise to four market segments. We define domestic and foreign securities included in the institutional investor's choice set as domestic and foreign *institutional securities* (I, I^*) and the remaining securities that are only invested by retail investors as domestic and foreign *retail securities* (R, R^*). The $N = N_R + N_I + N_{I^*} + N_{R^*}$ risky securities are partitioned into four segments: N_R domestic retail securities, N_I domestic institutional securities, N_{I^*} foreign institutional securities and N_{R^*} foreign retail securities. To simplify the notation, we omit time subscripts. We use column vectors $\mathbf{r}_R$, $\mathbf{r}_I$, $\mathbf{r}_{I^*}$ and $\mathbf{r}_{R^*}$ to denote the excess returns of domestic retail securities, domestic institutional securities, foreign institutional securities and foreign retail securities. For example, the $N_I \times 1$ vector of excess returns on domestic institutional securities is $\mathbf{r}_I = [r_{I_1}, \cdots, r_{I_{N_I}}]'$. The excess returns of securities at $t+1$ are joint normally distributed. Like many CAPM-type models, we also assume that the supply of risky securities is exogenous in terms of their market capitalization.[11] We also assume that the covariance structure between risky security returns is exogenous. We use a $n_I \times n_R$ matrix V_{IR} to denote the covariance

[10] Our choice of CARA-normal setup follows Errunza and Losq (1985).
[11] See Zerbib (2022) for a recent example.

matrix between domestic institutional and domestic retail securities, V_{jI} to denote a $1 \times n_I$ vector containing the covariance between domestic security j and domestic institutional securities.[12] Security risk premia denoted as μ are endogenously determined in the equilibrium by investors' optimization and market clearing conditions.

A5: Investors can borrow and lend at risk free rate r_f denominated in the reference currency. There is no short-sale constraint.

Our model features partial segmentation across country border as well as between institutional and retail segments. A distinct feature of our setup compared to those of international asset pricing models with partial segmentation (e.g. Errunza and Losq (1985), Chaieb and Errunza (2007)) is that there does not exist securities invested commonly by all investors. This is because our model distinguishes home-biased retail investors from global institutional investors. In particular, retail investors from different countries have non-overlapping investment scopes due to their home bias. As a consequence, local retail investors from different countries do not trade with each other directly. International risk-sharing is channeled by the institutional investor who invests in both countries. This makes our model similar to models featuring an arbitrageur trading in two segmented markets such as Gromb and Vayanos (2002) and Greenwood, Hanson, and Liao (2018). Unlike these models in which the arbitrageur or generalist trades all securities, the institution in our model only invests in a subset of securities in each country due to their mandate constraint. Moreover, unlike these models that have one security in each segmented market, we study the general case with multiple securities in each segment and do not impose strict restriction on their covariance structure. Our general setup nests these two types of existing models. When domestic and foreign institutional securities could perfectly substitute each other, the model collapses to the partial segmentation case discussed in Eun and Janakiramanan (1986) and Chaieb and Errunza (2007), with the addition of an institutional investor. When the institutional investor has no mandate con-

[12]We use similar notation for other covariance matrices and vectors.

straint and invests in all domestic and foreign securities, our setup is reduced to the market structure in Greenwood, Hanson, and Liao (2018).

B. Equilibrium risk premia

Because our model is symmetric between domestic and foreign countries, in what follows we focus on the equilibrium results for domestic investors and securities. The results in the foreign country are entirely symmetric hence not repeated.

We define the *domestic market portfolio* D as the value-weighted portfolio of all domestic securities. We use r_D to denote its excess return and M_D to denote its total market capitalization. We refer to the value-weighted portfolios of domestic institutional and of domestic retail securities as the domestic *institutional portfolio* I and the *domestic retail portfolio* R. We use r_I and r_R to denote the excess returns on domestic institutional and retail portfolios and M_I and M_R to denote their total market capitalization. The foreign market, institutional and retail portfolios are defined in the same way and their excess returns are denoted as r_F, r_{I^*} and r_{R^*} and their total market capitalization is denoted as M_F, M_{I^*} and M_{R^*}.

Given the set of institutional securities, we can decompose the return of any domestic investment j into a component that can be replicated using domestic institutional securities and a residual component by performing a multiple regression of its return r_j onto the returns of domestic institutional securities. We define security j's *attainable return* $\hat{r}_j$ as the return on the mimicking portfolio and the mimicking portfolio as the *attainable portfolio*:

$$\hat{r}_j = B_{jI}\mathbf{r}_I \tag{1}$$

where $B_{jI} = V_{jI}V_{II}^{-1}$ is a row vector of regression coefficients. The attainable exposure $\hat{r}_j$ represents the component of security j's return that can be attained by the institutional investor. The residual $\epsilon_j = r_j - \hat{r}_j$ is the component of security j that is only attainable by the domestic retail investor. By definition, the return of any institutional security is attainable. If a domestic retail security can be perfectly replicated by a portfolio of domestic institutional securities, then it is attainable although it is not directly invested by the institutional investor. We define the *attainable domestic market portfolio* $\hat{D}$ as the portfolio of domestic institutional securities that optimally mimics the domestic market portfolio and we use $r_{\hat{D}}$ to denote its excess return. The attainable domestic market portfolio is the value weighted portfolio of the domestic institutional portfolio and the attainable domestic retail portfolio:

$$r_{\hat{D}} = \frac{M_I}{M_D} r_I + \frac{M_R}{M_D} \hat{r}_R \tag{2}$$

where $\hat{r}_R$ is the attainable return on the domestic retail portfolio. The attainable foreign market portfolio $\hat{F}$ is defined in the same way. Due to its mandate constraint, the institutional investor does not invest in the domestic market portfolio but would invest in the attainable domestic market portfolio.

Because the institutional investor trades foreign institutional securities with home-biased foreign retail investor, international risk-sharing depends on how well the attainable domestic market portfolio can be substituted by foreign institutional securities. We further decompose the return on the attainable domestic market portfolio $\hat{D}$ into a component that can be replicated by a portfolio of foreign institutional securities and a residual component that is not attainable by the foreign retail investor. We define the *substitute portfolio* for the domestic market portfolio D^s as the portfolio of foreign institutional securities that optimally mimics the attainable domestic market portfolio and refer to its return as the *substitute return*. The substitute portfolio is constructed by running a multiple regression of the return on the attainable domestic market portfolio onto the returns of foreign institutional securities:

$$r_{D^s} = B_{\hat{D}I^*} \mathbf{r}_{I^*} \tag{3}$$

where $B_{\hat{D}I^*} = V_{\hat{D}I^*} V_{I^*I^*}^{-1}$ is a row vector of regression coefficients. In equilibrium, the institutional investor sells the substitute portfolio to the foreign retail investor, who is willing to hold it as their optimal substitute for domestic investment. Appendix B.3 provides further details about the equilbrium investment by each investor.

The following proposition provides the equilibrium risk premia of domestic securities[13].

PROPOSITION 1:

If A1-A5 are satisfied,

1. the equilibrium risk premium of any domestic securities j is:

$$\mu_j = \underbrace{\gamma M_W cov(\hat{r}_j, r_{\hat{W}})}_{\text{attainable world market premium}} + \underbrace{\frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(\hat{r}_j, r_{\hat{D}} - r_{D^s}\big)}_{\text{institutional local premium}} + \underbrace{\gamma^d M_R cov\big(r_j - \hat{r}_j, r_R - \hat{r}_R\big)}_{\text{retail local premium}} \quad (4)$$

2. particularly,

(i) the risk premium of any domestic institutional security I_j is:

$$\mu_{I_j} = \underbrace{\gamma M_W cov(r_{I_j}, r_{\hat{W}})}_{\text{attainable world market premium}} + \underbrace{\frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(r_{I_j}, r_{\hat{D}} - r_{D^s}\big)}_{\text{institutional local premium}} \quad (5)$$

(ii) the risk premium of any domestic retail security R_j is:

$$\mu_{R_j} = \underbrace{\gamma M_W cov(\hat{r}_{R_j}, r_{\hat{W}})}_{\text{attainable world market premium}} + \underbrace{\frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(\hat{r}_{R_j}, r_{\hat{D}} - r_{D^s}\big)}_{\text{institutional local premium}} + \underbrace{\gamma^d M_R cov\big(r_{R_j}, r_R - \hat{r}_R\big)}_{\text{retail local premium}} \quad (6)$$

where γ is the aggregate absolute risk aversion defined in B.31. $r_{\hat{W}}$ is the return on the attainable world market portfolio defined as the value-weighted portfolio of the attainable

[13]See Appendix B for proof.

portfolios from each country.

$$r_{\hat{W}} = \frac{M_D}{M_W} r_{\hat{D}} + \frac{M_F}{M_W} r_{\hat{F}}$$

$$M_W = M_D + M_F$$

Domestic securities earn three risk premia. We define the first risk premium as the *attainable world market risk premium* that compensates for the covariance between the attainable return of domestic security j and the attainable world market return. The attainable world market factor represents the component of the world market risk that is attainable to the global institutional investor. γ is the aggregate absolute risk-aversion of the economy. The aggregate risk aversion increases with the absolute correlation between domestic and foreign institutional securities because the equilibrium is solved from the first-order condition of the institutional investor.[14]

In addition to the attainable world market risk premium, domestic securities also earn two local risk premia. We define the first local risk premium as the *institutional local premium*. It is compensation for the covariance between the attainable exposure of domestic security j and the institutional local risk factor, defined as the return difference between the attainable domestic market portfolio and its substitute portfolio:

$$f^{ilocal} = r_{\hat{D}} - r_{D^s} \tag{7}$$

This first local risk premium arises because the substitute portfolio does not perfectly replicate the attainable domestic market portfolio. Hence the institution cannot perfectly hedge its position in one country by trading institutional securities from the other country. The price of the institutional local risk increases with the risk-aversion of the institutional investor γ^i and decreases with the risk aversion of the foreign retail investor γ^f. Intuitively, when the

[14]See Appendix B for a detailed discussion

institutional investor is more risk averse relative to the foreign retail investor, fewer foreign institutional securities are held by the institutional investor, reducing international risk-sharing from domestic to foreign and increasing domestic institutional local premium. The attainable world premium and the institutional local premium depends on the attainable exposure of a security. This is because these two risk premia are associated with cross-border risk sharing and any covariance between securities from different countries must be transmitted through their attainable exposure that is accessible to the institutional investor.[15] We define the second local risk premium as the *retail local risk premium*, which is compensation for the covariance between the residual in security j's return that is not attainable and the residual in the domestic retail portfolio's return that is not attainable, which we define as the retail local factor:

$$f^{rlocal} = r_R - \hat{r}_R \tag{8}$$

This risk premium is compensation for the component of security j's covariance with the domestic retail portfolio that cannot be explained by domestic institutional securities. The price of the retail local risk depends on the risk aversion of the domestic retail investor γ^d since this residual risk is held exclusively by the domestic retail investor. If j is a domestic institutional security, the retail local risk premium is zero because its return is fully attainable ($\hat{r}_j = r_j$).

Appendix B shows that our general pricing result (4) implies the following beta representation:

$$\mu_j = \beta^{\hat{W}} \mu^{\hat{W}} + \beta_j^{ilocal} \mu^{ilocal} + \beta_j^{rlocal} \mu^{rlocal} \tag{9}$$

[15] Karolyi and Wu (2018) have a similar indirect covariance term in their empirical specification but it is the indirect covariance through an investable set accessible to all investors. This is also the case in Chaieb and Errunza (2007), where the exposure to currency risk is also measured with respect to the mimicking portfolio for investable securities. In our model, the indirect correlation is through institutional securities within each country.

where $\mu^{\hat{W}}$, μ^{ilocal} and μ^{rlocal} are the risk premia of the attainable world market factor, the institutional local factor and the retail local factor and βs are defined in (B.47), (B.48) and (B.49).

The relative importance of global versus local risk premia depends on three aspects of global institutional investment. First, the coverage of the institution's mandate. If a security is covered in the institution's mandate or if its exposure can be spanned by institutional securities, its retail local risk premium would be zero. Second, the substitutability (correlation) between domestic and foreign institutional securities, which determines the quantity of risk that can be shared internationally through institutional investment. Third, the risk-bearing capacity of the institutional investor $1/\gamma^i$, which determines the level of the institutional local risk premium.

C. Testable implications

Our theory has the following testable implications. First, the institutional local risk premium should be positive in countries without a lot of foreign retail investment and the retail local risk premium should be positive in countries where institutional securities could not fully span retail securities. Second, although the model is static in nature, the model predicts that the institutional local premium increases when institutional investors' risk aversion increases relative to retail investors. The time-series implication of this prediction is that the institutional local premium should be higher when institutional investors are more risk averse relative to retail investors. This is the case when the risk-bearing capacity of financial institutions is reduced due to financial rather than fundamental shocks.[16] Third, in the cross-section, stocks with higher institutional ownership and stocks that are more correlated with local institutional securities should earn higher attainable world premium, higher institutional local premium and lower retail local premium. The price of the attainable world market factor incorporates the

[16] Akbari, Carrieri, and Malkhozov (2022) document when global institutions have less capacity to hold global securities due to tightening financial constraints.

risk-bearing capacity of all investors and the price of the institutional local factor incorporates the risk-bearing capacity of institutional and retail investors in a country. In contrast, the retail local risk premium reflects the the risk-bearing capacity of local retail investors only. Because average risk aversion decreases with larger investor base and better risk-sharing, we expect the price of the retail local factor to be higher due to lack of risk-sharing. The net effect is that, when a firm has higher institutional ownership or is more correlated with local institutional securities, it should have lower cost of capital.[17]

III. Empirical Framework

This section introduces our empirical framework. We first discuss how we construct the institutional and retail local factors, then we explain our econometric specification and identification technique to estimate the risk premiums of individual stocks, finally we describe the data sources used in this study. In the empirical sections, when it is clear from the context, we do not use bold font to distinguish a vector from a scalar.

A. Constructing asset pricing factors

We first construct three pricing factors predicted by the theory: the attainable world market factor, the institutional local factor, and the retail local factor according to their definition in Section II.B.

For each domestic country in our sample D, we construct the returns on the attainable market portfolio $r_{\hat{D}}$ and on the attainable retail portfolio $\hat{r}_R$ by regressing the returns on its market portfolio r_D and on its retail portfolio r_R onto the return on its institutional portfolio

[17]Appendix B.4 provides a detailed analysis of how the risk premium of a firm changes after it is included in the mandate of institutional investors.

r_I:

$$r_{\hat{D},t} = \beta_{\hat{D}I,t} r_{I,t}$$

$$\hat{r}_{R,t} = \beta_{RI,t} r_{I,t}$$

where $\beta_{\hat{R}I,t}$ and $\beta_{\hat{D}I,t}$ are coefficients from 36-month rolling regressions.[18]

The retail local factor is the return difference between the domestic retail portfolio and its attainable component :[19]

$$f_t^{rlocal} = r_{R,t} - \hat{r}_{R,t}$$

The institutional local factor is the component of the attainable domestic market portfolio that cannot be explained by the foreign institutional portfolio I^*:

$$f_t^{ilocal} = r_{\hat{D},t} - \beta_{\hat{D}I^*,t} r_{I^*,t}$$

where $\beta_{\hat{D}I^*,t}$ is the coefficient from a 36-month rolling regression.

Lastly, we construct the attainable world market factor as the value-weighted portfolio of attainable market portfolios from each country. Figure 1 compares the Net Asset Value (NAV) evolution of a one-dollar investment at the beginning of the sample period in the value-weighted portfolio of all stocks (henceforth the world market portfolio) and the attainable world market portfolio. The two portfolios have the same trend, which is driven by the common component in both world market portfolios, institutional stocks, which have larger market capitalization. Our theory predicts that the attainable world market portfolio repre-

[18] We use institutional stocks identified in January 2000 to construct the institutional portfolios for the rolling windows needed before 2000 when ownership data is not available.

[19] We acknowledge that our construction of country-level asset pricing factors involves separate estimation of time-varying correlation and conditional volatilities, which is not internally consistent with the restrictions on covariance structure we impose on the conditional estimation. This is a choice for simplicity.

sents world market risk that is shared through institutional investors. Notably, the cumulative return of the attainable world portfolio is higher than that of the world market portfolio, which suggests that on average, institutional investors invest in stocks with better performance.

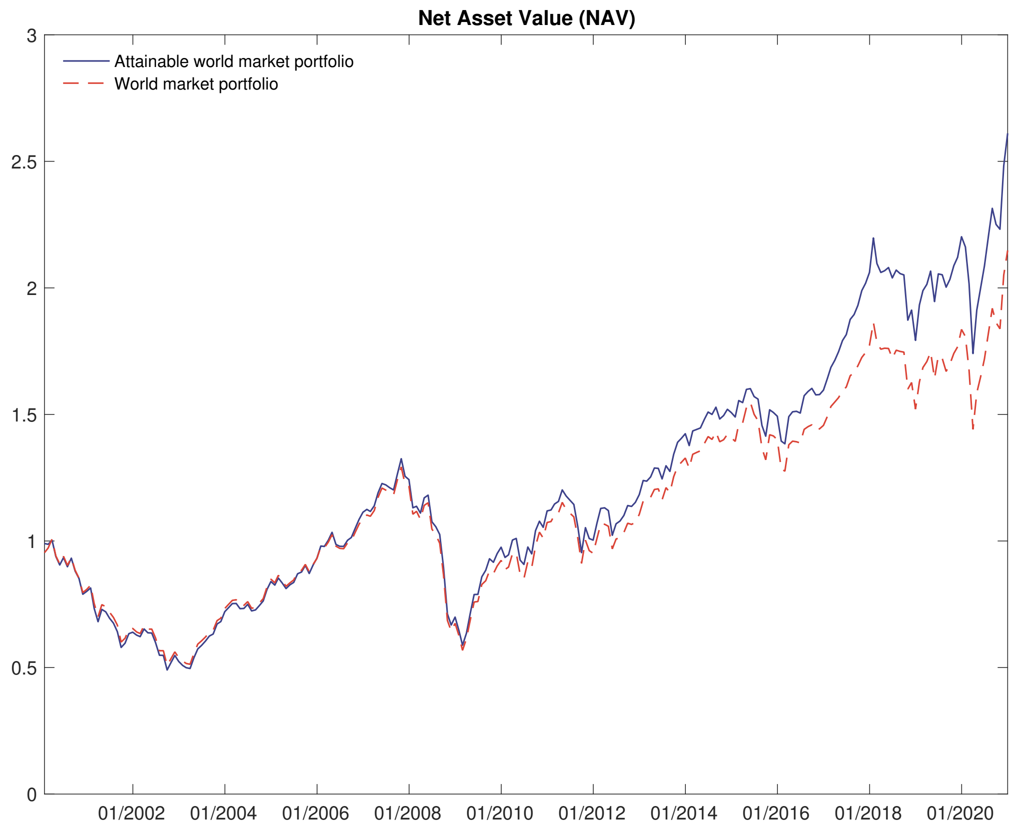

Figure 1. World market portfolio and attainable world market portfolio

This figure compares the world market portfolio and the attainable world market portfolio. It plots the evolution of the NAV over time of a one dollar investment in the world market portfolio and the attainable world market portfolio at the start of the sample period in January 2000.

B. Econometric specification

Our objective is to study how institutional ownership affects global and local risk premia. We would like to keep as much meaningful variation in institutional ownership in the cross section as possible given the small number of stocks in many of our sample countries and the limited coverage of ownership data. For this purpose, we estimate the model using individual stock returns data.[20]

We perform two estimations. We first perform a simple unconditional estimation of the covariance formulation of our pricing result (4) and estimate the price of covariance risk. We estimate the following empirical specification for the covariance formulation of our equilibrium pricing result (4):

$$E[r_i] = \alpha_i + \lambda^{\hat{W}} cov(r_i, r_{\hat{W}}) + \lambda^{ilocal} cov(r_i, f^{ilocal}) + \lambda^{rlocal} cov(r_i, f^{rlocal}) \qquad (10)$$

where $\lambda^{\hat{W}}$, λ^{ilocal} and λ^{rlocal} are the prices of the attainable world market, institutional local and retail local covariance risk. We perform a two-pass cross-sectional regression (Fama and MacBeth, 1973) with Newey and West (1987) standard errors to account for heteroskedasticity and serial correlation. In the first pass, we calculate for each stock three covariances using a 36-month rolling regression. In the second pass, for each month, we run cross-sectional regression of returns on a constant and the covariances estimated from the first pass. The estimated prices of risk are the average over cross-sections.

To study the time-variation in institutional and retail local risk premia, we also perform a conditional estimation of the beta representation of our main pricing result (9). We adopt a two-pass regression technique developed by Gagliardini, Ossola, and Scaillet (2016) and adapted by Chaieb, Langlois, and Scaillet (2021) for international setting. This technique extends the classic two-pass regression to the conditional setting and specifies time-varying factor exposure and factor risk premia as linear functions of lagged common and stock-specific instruments. The method is suitable for individual stocks and corrects for the error-in-variable problem from first-pass estimation.

The conditional version of the beta representation of our main pricing equation (9) for

[20]As is explained in Appendix C, to limit the effect of outliers, we winsorize returns in each country-month at the 1% and 99% level.

security i in country c implies that:[21]

$$E_{t-1}[r_{i,t}] = \beta'_{i,t}\mu_{c,t} \tag{11}$$

where $\mu_{c,t} = [\mu_t^{\hat{W}}, \mu_{c,t}^{ilocal}, \mu_{c,t}^{rlocal}]$ is a vector of conditional factor risk premia and $\beta_{i,t} = [\beta_{i,t}^{\hat{W}}, \beta_{i,t}^{ilocal}, \beta_{i,t}^{rlocal}]'$ is a vector of factor loadings of stock i.

We assume that the excess return of security i in country c has the following linear factor structure:

$$r_{i,t} = \alpha_{i,t} + \beta'_{i,t} f_{c,t} + \epsilon_{i,t} \tag{12}$$

where $f_{c,t} = [f_t^{\hat{W}}, f_{c,t}^{ilocal}, f_{c,t}^{rlocal}]'$ is a vector of factors constructed in Section III.A, $\beta_{i,t} = [\beta_{i,t}^{\hat{W}}, \beta_{i,t}^{ilocal}, \beta_{i,t}^{rlocal}]'$ is a vector of factor loadings of stock i.

The conditional pricing equation (11) and the linear empirical specification (12) imply the following asset pricing restriction:

$$\alpha_{i,t} = \beta'_{i,t}\nu_{c,t} \tag{13}$$

where $\nu_{c,t} = \mu_{c,t} - E_{t-1}[f_{c,t}] = (\Lambda_c - F_c)Z_{c,t-1}$ measures the wedge between factor risk premia and the conditional expected returns of factor portfolios. Although our asset pricing factors are tradable in theory, in practice, their implementation faces transaction costs due to rebalancing and short-selling, a non-zero ν captures these market imperfections (Gagliardini, Ossola, and Scaillet, 2016).

We incorporate conditional information using common and firm-specific instruments. First, factor loadings $\beta_{i,t}$ is a linear function of conditioning information at time $t-1$, which includes p lagged instruments $Z_{c,t-1}$ that are common to all stocks within a country c and q

[21] We acknowledge that our model is static in nature, a full-fledged conditional model would induce additional risk premia for hedging changes in investment opportunities, which is beyond the scope of this paper.

lagged instruments $Z_{i,t-1}$ that are stock-specific:

$$\beta_{i,t} = B_i Z_{c,t-1} + C_i Z_{i,t-1} \tag{14}$$

where B_i is a $3 \times p$ matrix and C_i is a $3 \times q$ matrix.

Second, we specify factor risk premia in country c, $\mu_{c,t}$ as a linear function of common instruments:

$$\mu_{c,t} = \Lambda_c Z_{c,t-1} \tag{15}$$

Third, we specify the conditional expectation of factors as a linear function of lagged common instruments:

$$E_{t-1}[f_{c,t}] = F_c Z_{c,t-1} \tag{16}$$

Our asset pricing restriction (13) together with our empirical specifications (14), (15) and (16) imply that:

$$\alpha_{i,t} = Z'_{c,t-1} B'_i (\Lambda_c - F_c) Z_{c,t-1} + Z'_{i,t-1} C'_i (\Lambda_c - F_c) Z_{c,t-1} \tag{17}$$

Thus the intercept $\alpha_{i,t}$ in the linear specification (12) can be expressed as a quadratic function of lagged instruments. $\beta'_{i,t} f_{c,t}$ are factors scaled by common instruments $Z_{c,t-1}$ and stock-specific instruments $Z_{i,t-1}$ under (14). We estimate (12) in the first pass by regressing $r_{i,t}$ on these two groups of regressors from $\alpha_{i,t}$ and $\beta_{i,t} f_{c,t}$ respectively:

$$r_{i,t} = b'_{1,i} x_{1,i,t} + b'_{2,i} x_{2,i,t} + \epsilon_{i,t} \tag{18}$$

where the first group $x_{1,i,t}$ contains all interaction terms among instruments from the quadratic form in $\alpha_{i,t}$ and the second group $x_{2,i,t}$ contains all factors scaled by instruments from $\beta'_{i,t} f_{c,t}$. Detailed definition of each regressor is provided in Appendix D.1. From the first-step estimate for $\hat{b}_{2,i}$, we can calculate estimates for coefficients in the time-varying beta specification $\hat{B}_i$ and $\hat{C}_i$ using the relation (D.9) and time-varying beta exposure $\hat{\beta}_{i,t}$ of each individual stock through (14).

Specifically, we use as common instruments a constant, the world dividend yield and the country dividend yield. $Z_{c,t-1} = [1, DY_{t-1}, DY_{c,t-1}]$ so that $p = 3$. The dividend yields are standardized to have zero mean and unit standard deviation. As stock instruments, we use the country-level percentile rank of a stock's size so that $q = 1$.[22] As is explained in Appendix D.1, using all instruments leads to a total number of 21 regressors in the time-series regression. In the data, the sample size of asset i can be small, resulting in unreliable estimates of $\hat{b}_i = (b'_{1,i}, b'_{2,i})'$. Therefore GOS apply trimming conditions to select stocks from the first-pass regression that have more than 60 months of monthly observations and have time series regression that is not too badly conditioned. Chaieb, Langlois, and Scaillet (2021) (CLS) show that applying the same trimming condition on international data results in a few or even zero stocks kept for several countries. To keep more stocks in the sample, they introduce an automatic selection procedure to select common instruments for each stock to reduce the number of regressors. We extend their procedure and impose selection also on stock-specific instruments. Unlike CLS who require that at least one stock-specific instrument be kept for a stock to be included in their sample, we drop the stock-specific instrument if it does not help explaining the time-variation of the stock's factor exposure. Details about our instrument selection procedure is provided in Appendix D.2.

The second-pass regression consists of running a cross-sectional weighted least squares

[22] We do not use institutional ownership as an instrument because institutional ownership (IO) is persistent in the data, using it will induce multicollinarity in the first-pass regression. For example, if the IO of stock i is relatively constant, then the interaction terms $IO_{i,t-1} f_{c,t}$ is highly correlated with $f_{c,t}$.

regression of $\hat{b}_{1,i}$ onto $\hat{b}_{3,i}$: $b_{1,i} = b_{3,i}\nu_c$, where $\hat{b}_3$ is a transformation of $\hat{b}_2$ defined in (D.4), and ν_c is the vectorized form of $\Lambda_c - F_c$ defined in (D.5). Finally, we estimate F_c by running a seemly unrelated regression (SUR) of $f_{c,t}$ on the common instruments $Z_{c,t-1}$ and obtain Λ_c through the relation $\nu_c = vec(\Lambda_c' - F_c')$. The loading of time-varying risk premia $\mu_{c,t}$ on common instruments Λ_c in (15) has the following components:

$$\Lambda_c = \begin{bmatrix} \Lambda_0^{\hat{W}} & \Lambda_{DY}^{\hat{W}} & \Lambda_{DY_c}^{\hat{W}} \\ \Lambda_0^{ilocal} & \Lambda_{DY}^{ilocal} & \Lambda_{DY_c}^{ilocal} \\ \Lambda_0^{rlocal} & \Lambda_{DY}^{rlocal} & \Lambda_{DY_c}^{rlocal} \end{bmatrix} \tag{19}$$

where Λ_0, Λ_{DY} and $\Lambda_{DY,c}$ are the loadings of time-varying risk premia on the constant, world dividend yield and country dividend yield. Because we standardize world and country dividend yields to have zero mean and unit standard deviation, the value and significance of Λ_0s are the levels and significance of the unconditional risk premia. We compute country-level time-varying risk premia as $\hat{\mu}_{c,t} = \hat{\Lambda}_c Z_{c,t-1}$. Using time-varying factor loadings $\hat{\beta}_{i,t}$ estimated from the first-pass and risk premia estimate $\hat{\mu}_{c,t}$ from the second-pass, we calculate stock-level model-predicted risk premia as $\mu_{i,t} = \beta_{i,t}'\mu_{c,t}$. Note that because $\beta_{i,t}'$ is estimated from the first-pass regression, we can calculate time-varying risk premia for all stocks included in the first-pass regression that have more than 60 months of observation even if they are excluded from the second-pass regression. We delegate further details about the estimation procedure and inference to Appendix D.

We impose two restrictions on our estimation following Chaieb, Langlois, and Scaillet (2021). First we require that the loading of the conditional expectation of the attainable world factor on country dividend yield $F_{DY_c}^{\hat{W}}$ is zero. This condition ensures that the world factor has the same conditional expectation across countries. Second, for the sake of parsimony, we impose the restrictions that factor loadings $\beta_{i,t}$, risk premia $\nu_{c,t}$ and the conditional expectations of local factors $E_{t-1}[f_{c,t}^{ilocal}, f_{c,t}^{rlocal}]$ do not load on the global instrument. These restrictions have two implications. First, local risk premiums do not load on the global instrument $\Lambda_{DY}^{ilocal} = \Lambda_{DY}^{rlocal} = 0$. Second, the loading of the global risk premium on the global instrument is the same as the loading of the conditional expectation of the global factor on the global instrument $\Lambda_{DY}^{\hat{W}} = F_{DY}^{\hat{W}}$. Thus $\Lambda_{DY}^{\hat{W}}$ is the same across countries and the global risk premium depends on the global instrument only through its conditional expectation not through the risk premium ν_c.

C. Data

We study 38 countries including 23 DMs and 15 EMs.[23] The sample period is from January 2000 to December 2020 with monthly frequency. We calculate monthly USD returns, applying standard filters and data corrections ending up with 67,049 individual stocks from the Compustat Global universe.[24] USD risk-free rates are calculated using the one-month T-bill rate from WRDS Fama-French daily research factors. We obtain dividend yield at the country and world level from Datastream. Quarterly security-level institutional ownership ratio is calculated from 2000 Q1 to 2020 Q4 using the FactSet ownership database following Ferreira and Matos (2008). Because our focus is on the international investment of institutional investors, we focus on global institutional investors who do not invest only in one country or region. For each quarter, we calculate the country and region weight in the portfolio of each institution. Following Bartram et al. (2015), we classify an institution as a global institution if the maximum percentage of its holdings in a country does not exceed 90% and the maximum percentage of its portfolio in a region does not exceed 80%. We merge the Compustat universe and FactSet using commonly used identifiers (CUSIP, ISIN, SEDOL). We assign zero institutional ownership to a Compustat observation that cannot be matched with

[23]Our selection criterion is that these countries are included in the FTSE All Country World Index, have FactSet coverage since 2000 and have at least 10 stocks with positive global institutional ownership by the end of 2003.

[24]See the online appendix of Chaieb, Langlois, and Scaillet (2021) for a very detailed analysis of available sources for international equity data.

FactSet. For each month, we use the ownership ratio that is available in the corresponding quarter. Details about data construction is provided in Appendix C.

We use firm-level ownership by global institutions to classify institutional and retail securities. A security is classified as an institutional security if its global institutional ownership exceeds the 50^{th} percentile in its country in the quarter and is higher than 1%. For each domestic country D, we construct the domestic market portfolio D as the value-weighted portfolio of all stocks issued by firms that are domiciled in the country, the institutional portfolio I as the value-weighted portfolio of all domestic institutional stocks and the retail portfolio R as the value-weighted portfolio of all domestic retail stocks. In addition, we construct foreign institutional portfolio I^* as the value-weighted portfolio of institutional stocks outside of the country. Figure 2 shows a heat map of the pairwise correlation between institutional portfolios and between retail portfolios from different countries. The upper-triangular half of the figure presents the correlation between institutional portfolios from each pair of countries, the lower-triangular half shows the correlation between retail portfolios. Countries are ordered by their region: Latin America, North America, Europe, Pacific developed, EMEA, and Asian emerging. Overall, high correlation is concentrated in the block among European and North American countries. The correlation between these markets and other markets and the correlation among other markets are lower. In addition, the figure shows that the correlations between institutional portfolios in the upper-triangular half are higher than the correlations between retail portfolios in the lower-triangular half. The lower correlation between retail segments across countries is a reflection of the fact that these remote markets do not share a lot of common investors.[25]

[25] Although the correlation structure is exogenous in our model, ownership by common investors could increase the correlation between stocks (Anton and Polk, 2014).

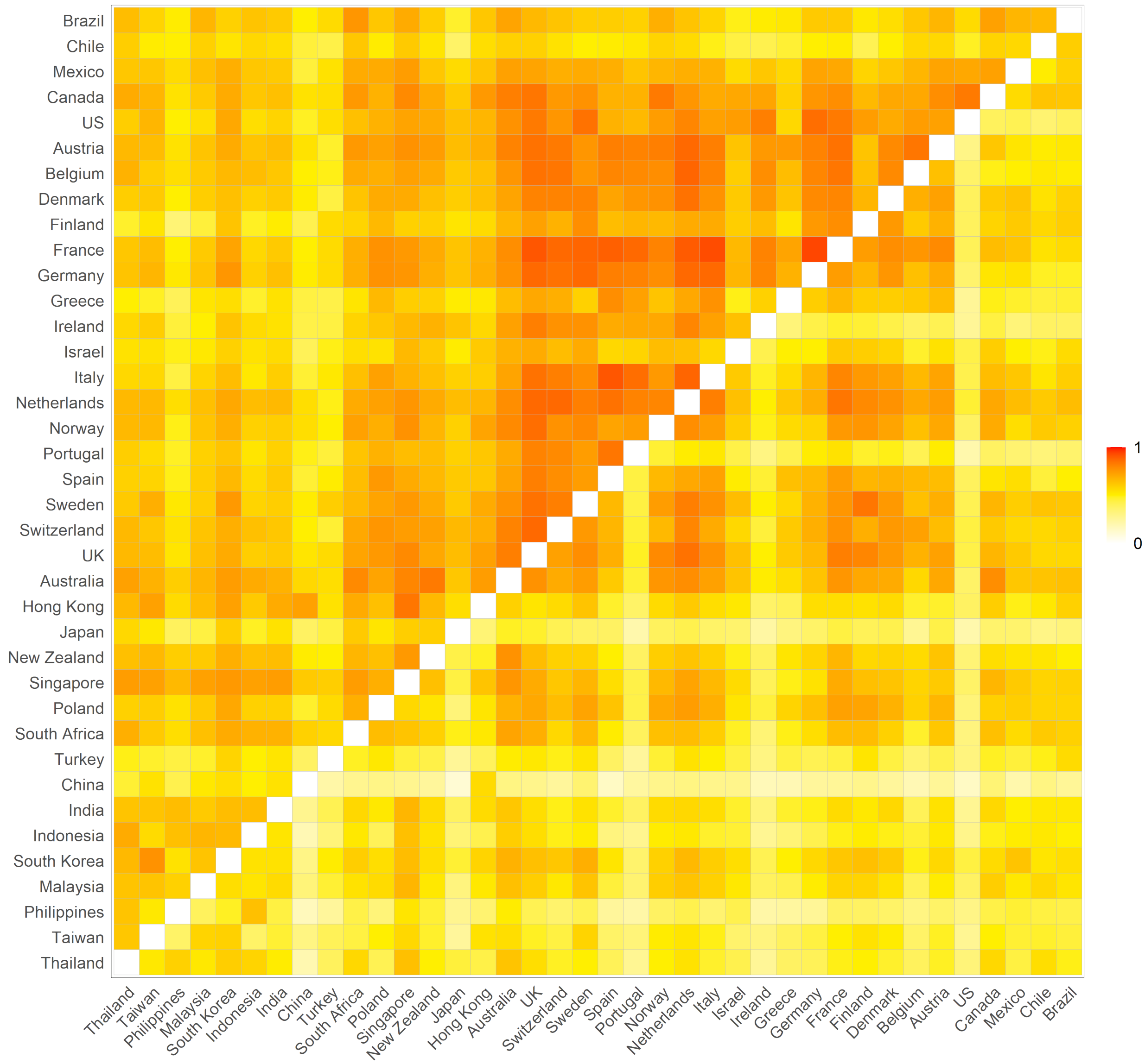

Figure 2. Pairwise correlation between institutional portfolios and between retail portfolios across countries

This figure shows the pairwise correlation between institutional and retail portfolios from different countries in our sample from January 2000 to December 2020. The upper triangular half of the matrix are the pairwise correlations between institutional portfolios from two countries. The lower triangular half of the matrix are the pairwise correlations between retail portfolios from two countries.

We construct the institutional local factor and the retail local factor using rolling regressions as explained in Section III.A. Table I reports for each country the total number of stocks in our data, the average proportion of institutional stocks to the total market capitalization of each country and the annualized average returns and volatilites of the institutional local factor and the retail local factor. In most DMs, institutional stocks represent more than 80% of total market capitalization. In US for example, the share of institutional securities in total market capitalization is as high as 91.1%. This average proportion is higher in DMs than in EMs because at the beginning of our sample, less than 50% of firms have global institutional ownership that is higher than 1% in many EMs. The average return of the institutional local factor is positive in 13 of 23 DMs and is positive in 12 of 15 EMs in our sample. On the other hand, the retail local factor is negative in 16 DMs and 14 EMs. Because the retail local factor is constructed as the residual return in the retail portfolio that cannot be explained by the institutional portfolio, its negative average return suggests that this residual component in retail securities performs badly in many countries during our sample period. This is consistent with what we show in Figure 1 that the world market portfolio underperforms the attainable world market portfolio.

IV. Empirical Results

In this section, we present and analyze our empirical results. First, we discuss the significance and level of the attainable world market premium, the institutional local premium and the retail local premium across countries. Second, we analyze how institutional and retail local risk premia vary over time in DMs and EMs. Third, we study how firm-level institutional ownership affects global and local risk premia.

A. Significance and level of global and local risk premia across countries

Table II presents the results of our unconditional estimation (10). We expect that the price of the institutional local risk be positive and significant in markets without a lot of foreign retail investment and have institutional securities that are not highly correlated with foreign institutional securities. We expect the price of the retail local risk to be positive and significant in markets where institutional securities could not span retail securities. In

US, which is a benchmark for the most open and integrated market, only the price of the attainable world market risk is positive and significant. In other markets, the price of the institutional local risk is positive and significant at the 5% level or less in 15 out of 23 DMs. The price of the retail local risk is positive and significant in 7 DMs. Both institutional and retail local risk factors are significantly priced in Austria, Canada and Israel, meaning that the two local factors explain different components of the cross-section of individual stock returns in these countries. In 10 DMs, only the institutional local risk is positively and significantly priced, indicating that although institutional securities in these countries could not be spanned by foreign institutional securities, institutional securities could span retail securities and risk-sharing between institutional investors and retail investors is perfect. In Denmark, Norway and Singapore, only the retail local factor is significantly priced, meaning that although institutional securities are priced globally, the retail local factor is needed to explain the returns of retail securities. Overall, we find that institutional and retail factors are significantly priced across a wide range of DMs. We also find strong evidence for the significance of the two local factors in EMs. The institutional local factor is positively and significantly (at the 10% level or less) priced in 9 out of 15 emerging markets and the retail local factor is positively and significantly priced in 6 emerging markets. Both institutional local and retail local factors are significantly priced in China, Indonesia and Philippines. Therefore retail securities in these countries could enjoy better risk-sharing if institutional investors enlarge their investment mandate. Whereas in India, Malaysia, Mexico, South Korea, Taiwan and Turkey, only the institutional local factor is significant. In summary, we find wide-spread evidence that the institutional and retail local local factors are significantly and positively priced in the cross-section of individual stocks. We also reveal rich cross-country differences: in some countries, both institutional and retail local factors are significant. In other countries, only one of the local factors is significantly priced.

We estimate time-varying factor risk premia in our conditional estimation. In the conditional specification (15), time-varying risk premia is linear in common instruments. Table III presents our estimates of the loadings of the attainable world market risk premium, the institutional local risk premium and the retail local risk premium on the constant Λ_0 and on the local dividend yield Λ_{DY_c} in (15). As is explained in Section III.B, coefficients $\Lambda_0^{\hat{W}}$, Λ_0^{ilocal} and Λ_0^{rlocal} measure whether the unconditional risk premia of the attainable world market factor, the institutional local factor and the retail local factor are significant in the cross-section of individual stocks in each country.[26] Λ_{DY_c} measures whether the local premia vary significantly over time with local market conditions proxied by the local dividend yield. Because we impose the restriction that $\Lambda_{DY}^{ilocal} = \Lambda_{DY}^{rlocal} = 0$ and that $\Lambda_{DY}^{\hat{W}}$ is the same across all countries, we only report the coefficients on the local instrument Λ_{DY_c}. Because we do not impose any restriction on the sign of the estimated unconditional risk premia Λ_0s, due to estimation noise, some of the estimates can be negative. This is not a new finding in this paper. In their study of US individual stocks, GOS find that the value premium is negative and in their study of interntaional individual stocks, Chaieb, Langlois, and Scaillet (2021) (CLS) find that the average risk premia of the excess country factor (defined as the difference between the country's market return and the world market return) to be negative in DMs.

The unconditional attainable world market risk premium is significant and positive at the 5% level or below in 11 out of 23 DMs and 6 out of 15 EMs. The proportion of DMs and EMs with significant world risk premium is comparable but lower than CLS, who find 61% of DMs and 71% of EMs have significant world market risk premium using a larger set of 47 countries over a longer and earlier sample period from 1985 to 2018. The unconditional institutional local risk premium is positive in 15 out of 23 DMs. It is significantly positive at the 5% level or below in Belgium and Denmark and is significantly positive at the 10% level in Hong Kong and Switzerland. The unconditional retail local risk premium is positive

[26] In order to make valid inference based on the GOS approach, there should not be any remaining factor structure in the residuals of (12). We calculate the diagnostic criteria of Gagliardini, Ossola, and Scaillet (2019) and verify that the diagnostic criteria is negative in 34 countries except China, Greece, Ireland and Spain, meaning that there is no remaining factor structure in the residuals in most of our sample countries.

in 15 out of 23 DMs. It is significant at the 5% level in Austria, Ireland and Israel and at the 10% level in Hong Kong. In our unconditional estimation, we also find that the price of the institutional local risk is significant in Hong Kong, Denmark and Switzerland and the price of the retail local risk to be significant in Austria and Israel. However, we find overall less evidence for the statistical significance of the local risk premia in DMs although they are estimated to be positive in most of our sample countries. There is stronger evidence for the significance of institutional and retail local risk premia in EMs. The unconditional institutional local premium is positive in 12 out of 15 DMs. It is significant and positive at the 5% level or below in India, Indonesia, Philippines, South Africa and Thailand and is significant at the 10% level in Malaysia. The unconditional retail local premium is positive in 10 EMs. It is significant at the 5% level in Greece and is significant at the 1% level in Turkey. Across all 23 DMs, only 5 countries have one of the local risk premia that is positive and significant at the 5% level. 7 out of 15 EMs have at least one local risk premuim that is significantly positive. Compared to our unconditional test, fewer countries have significant average local risk premia in our conditional estimation. This is because in our conditional estimation, we apply triming conditions and keep only stocks that have a sufficiently long sample period (60 months) and well-conditioned first pass time-series regression, whereas in our unconditional estimation, we do not apply such triming conditions. Therefore our unconditional estimation keeps more stocks in the sample, as it requires only estimating covariance in the first-step using a rolling window, which allows us to keep more dispersion in the cross-section. The trimming condition could explain the lack of significance of the retail local premium. Exposure to the retail local factor should better explain the cross-section of retail securities that are outside of the investment mandate of global institutional investors. Retail stocks tend to be issued by firms that are younger and have shorter time-series, therefore we are likely to exclude more retail stocks from the second-pass regression. Table III also reports the number of stocks used in the second-pass cross-sectional regression to estimate factor risk premia.

For most countries, compared to the total number of stocks used to construct pricing factors in Table I, only about a half of all stocks in the Compustat universe are included in the second-pass estimation. This likely explains why we do not find significant retail local risk premium in most countries in the conditional estimation.

The proportion of countries with significant local risk premia in our conditional estimation is lower than the findings of CLS that the excess country market risk premium is significantly positive in 43% DMs and 54% EMs in the world four-factor model with world market, size, value and momentum factors. We do find weaker evidence for the significance of local factors compared to CLS because their sample is earlier and includes more less-developed countries. Their sample starts in the 1990s for most countries, whereas our more recent sample starts in 2000. International stock markets have become more integrated in the recent sample. They are also able to include more countries because they do not require the availability of sufficient number of institutional stocks to form the institutional portfolio in each country. The earlier sample period and the inclusion of markets with less global institutional investment could explain why they have more countries with significant local risk premium.

On average, the institutional local risk premium is important in both DMs and EMs compared to the attainable world market risk premium. The unconditional institutional local risk premium is on average 2.76% in DMs, which is about half the level of the average unconditional attainable world market premium of 5.51% . In EMs, the average institutional local risk premium is 6.27% , which is higher than the attainable world market risk premium of 4.23% . The average unconditional retail local risk premium is 1.71% in DMs. In EMs, the average retail local risk premium is 2.65% . The institutional local premium is lower in DMs for two reasons. First, our theory predicts that the institutional local risk premium of a country should be lower when its institutional securities can be better substituted by foreign institutional securities. Institutional securities in DMs are more correlated with institutional securities from foreign countries as is shown in Figure 2, which explains the lower institutional local risk premium in DMs. Second, our model predicts that the institutional local risk premium is lower when institutional investors investing in the country have higher risk-bearing capacity. We also plot unconditional institutional risk premium against country-level average global institutional ownership in Figure 3. The institutional local risk premium decreases with average country institutional ownership, which is a proxy for the risk-bearing capacity of institutional investors investing in the country.

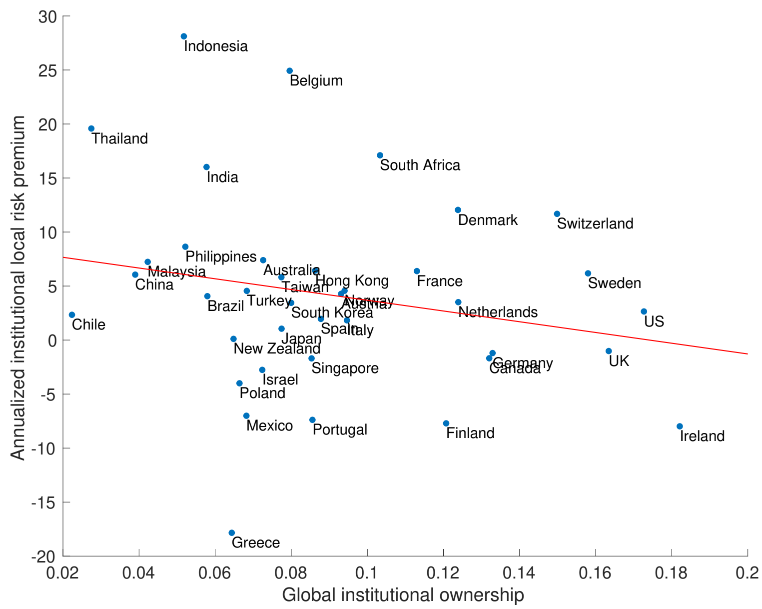

Figure 3. Unconditional institutional local risk premium and country level global institutional ownership

This figure plots the unconditional institutional local risk premium against average country-level global institutional ownership. The unconditional institutional local risk premium is estimated from our conditional estimation and reported in Table III. Country-level global institutional ownership is the time-series average of the market capitalization of each country that is owned by global institutional investors.

B. How do global and local risk premia vary over time?

The coefficient of DY_c reported in Table III indicates whether global and local risk premia vary significantly over time with the country dividend yield. Country dividend yield has been widely used in existing studies as an instrument for conditional information. It is a proxy for local market conditions. The country dividend yield significantly explains time-variation in the attainable world premium at a significance level of 5% or below in 8 out of 23 DMs and 5 out of 15 EMs. In DMs, the institutional local premium only varies significantly with the local dividend yield in Belgium and Portugal but the retail local risk premium varies significantly with the local dividend yield in 4 DMs. Local dividend yield significantly explains the time variation in the institutional local risk premium in 5 EMs and the time-variation in the retail local premium in 3 EMs.

Figure 4 plots the value-weighted average of time-varying annualized attainable world market premium, institutional local premium and retail local premium across DMs. In each month, we calculate value-weighted average using the aggregate market capitalization of each country. In DMs, both the attainable world market premium and the institutional local premium are important drivers of time-varying total risk premium. These two risk premia spiked during distressed episodes including the Global Financial Crisis (GFC) and the Covid crisis. In contrast, there is little variation in the retail local risk pemium in DMs.

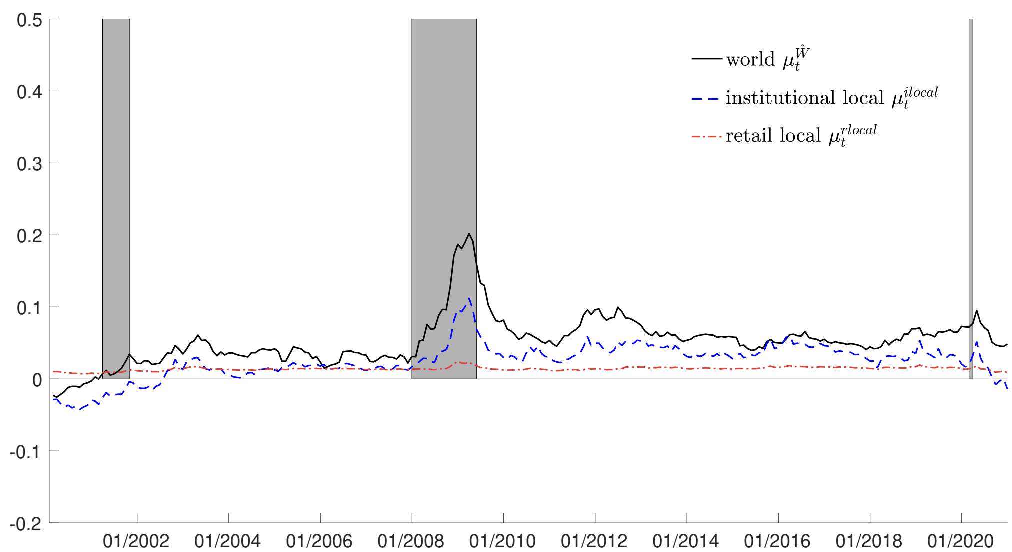

Figure 4. Annualized risk premia in developed markets

This figure plots the value-weighted average across countries of annualized estimated time-varying attainable world market premium $\mu_{c,t}^{\hat{W}}$, institutional local premium $\mu_{c,t}^{ilocal}$ and retail local premium $\mu_{c,t}^{rlocal}$ in developed markets (DMs). For each country, the time-varying risk premium is calculated using estimated coefficients Λs reported in Table III and time-varying common instruments through (15). We report the average weighted by total market capitalization of each country.

Figure 5 plots the value-weighted time-varying risk premia in EMs. Our estimation reveals interesting time-series dynamics in EMs. First, the attainable world market premium is higher in the second half of the sample than in the first half. This is consistent with better risk-sharing in EMs due to the increase in the level of global institutional investment over time. Second, the retail local premium explains more of the time-variation of total risk premium in EMs. Interestingly, the institutional and retail local risk premia exhibit different dynamics during distressed episodes. During both the GFC and the Covid crisis, the institutional local risk premium spiked. This finding is consistent with the empirical evidence that institutional investors, especially foreign institutional investors, tend to reduce their average stock investments during periods of global market stress (Kacperczyk, Nosal, and Wang, 2022). This could be due to institutional investors being subject to tighter financial constraints and having reduced risk-bearing capacity (Akbari, Carrieri, and Malkhozov, 2022). In contrast, the retail local premium only increased during the Covid crisis, not the GFC. The convergence of institutional and retail local risk premia during the Covid crisis and the divergence of the two risk premia during the GFC show that the former episode arises from fundamental shocks and the later arises from financial shocks.

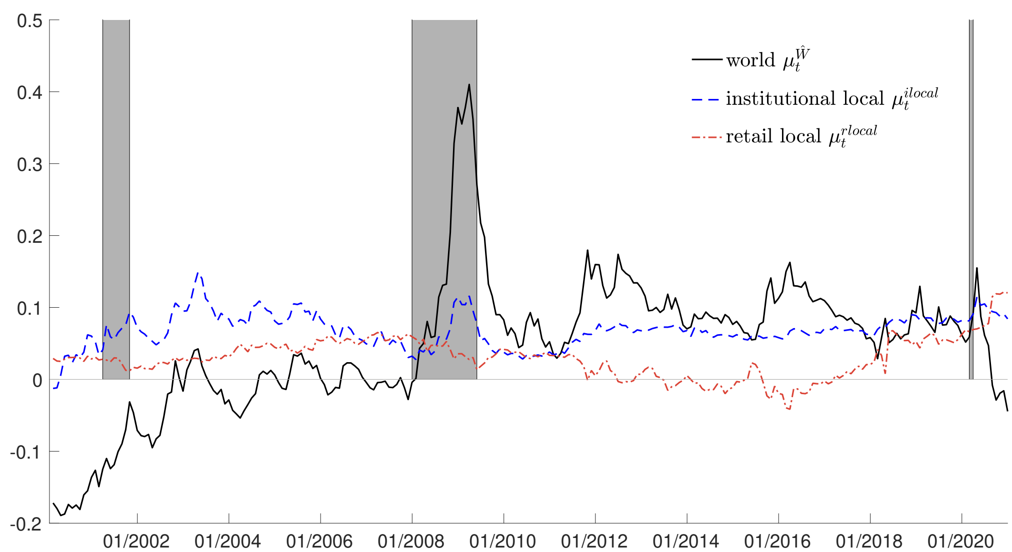

Figure 5. Annualized risk premium in emerging markets

This figure plots the value-weighted average across countries of annualized estimated time-varying attainable world market premium $\mu_{c,t}^{\hat{W}}$, institutional local premium $\mu_{c,t}^{ilocal}$ and retail local premium $\mu_{c,t}^{rlocal}$ in emerging markets (EMs). For each country, the time-varying risk premium is calculated using estimated coefficients Λs reported in Table III and time-varying common instruments through (15). We report the average weighted by total market capitalization of each country.

C. How does institutional ownership affect local and global risk premia in the cross section?

We discussed in Section IV.A and Section IV.B how global and local risk premia predicted by our model vary across countries and over time. In this section, we would like to quantify how global institutional ownership affects the level of global and local risk premia in the cross-section of individual stocks. As is explained in Section III.B, from our conditional estimation, we calculate model-implied time-varying total risk premia through $\mu_{i,t} = \beta'_{i,t}\mu_{c,t}$ for each individual stock that has more than 60 months of observation. We could then analyze how stock-level risk premium and its global and local components vary with its global institutional ownership. Table IV reports the average composition of the model-implied risk premia of individual stocks. We first calculate for each month the equal-weighted average of model-implied risk premia across individual stocks, then report the time-series average for each country. The average risk premium across individual stocks is similar to the country-level unconditional risk premium reported in Table III.

To study how institutional ownership affects local and global risk premia, we regress model-predicted risk premia on institutional ownership and firm-level, country-level and time-varying controls. Specifically, we consider as regressors the total risk premium $\mu_{i,t} = \beta'_{i,t}\mu_{c,t}$, the attainable world market premium $\mu^{\hat{W}}_{i,t} = \beta^{\hat{W}}_{i,t}$, the institutional local premium $\mu^{ilocal}_{i,t} = \beta^{ilocal}_{i,t}\mu^{ilocal}_{c,t}$ and the retail local premium $\mu^{rlocal}_{i,t} = \beta^{rlocal}_{i,t}\mu^{rlocal}_{c,t}$. We run the following panel regression:

$$
\begin{aligned}
y_{i,t} =& \beta_1 IO_{i,t-1} + \beta_2\rho_i + \beta_3 X_{i,t-1} + \beta_4 CountryIO_{c,t-1} + \beta_5 CR_{t-1} + \epsilon_{i,t} \\
& y \in \{\mu_{i,t}, \mu^{world}_{i,t}, \mu^{ilocal}_{i,t}, \mu^{rlocal}_{i,t}\}
\end{aligned} \tag{20}
$$

where $IO_{i,t-1}$ is the lagged firm-level institutional ownership, ρ_i is the time-invariant correlation between security i and the institutional portfolio of its country, $X_{i,t-1}$ are lagged firm level controls including log market capitalization (logmv), book-to-market ratio (bm) and dividend yield (dy). Because our model predicts that the institutional local risk premium depends on the risk-bearing capacity of global institutional investors, we also include two proxies for institutional risk-bearing capacity. $CountryIO_{c,t-1}$ is the lagged country-level global institutional ownership calculated as the value-weighted global institutional ownership across all stocks in country c. Country-level institutional ownership captures variation in institutional risk-bearing capacity across countries. CR_{t-1} is the lagged intermediary capital ratio of He, Kelly, and Manela (2017). The capital ratio captures variation in institutional risk-bearing capacity over time. To control for omitted variables that vary with country and time, we also report an alternative specification in which we include country-time fixed effects. Our theory predicts that in the cross-section, stocks with higher institutional ownership and stocks that are more correlated with institutional stocks are more attainable hence should earn lower retail local risk premium and higher attainable world market premium.

Table V presents the results for DMs. The attainable world risk premium and the institutional local risk premium increase significantly with IO and ρ in both specifications. When controlling for country-month fixed effect, a 1% increase in firm-level IO predicts an increase of 1.4 bps in the attainable world market premium and an increase of 1.9 bps in the institutional local premium. This is consistent with the prediction of our theory that stocks with higher institutional ownership and stocks that are more correlated with institutional stocks are more attainable hence earn higher world and institutional local risk premia. The retail local premium decreases with firm-level IO in regression (7) with country and time controls but increases with IO in regression (8) with country-month fixed effect. The positively significant coefficient in (8) is small in terms of economic magnitude and can be explained by estimation noise in the risk premium from the conditional estimation because the retail local factor is not significantly priced in most developed markets. The net effect of institutional ownership on total risk premium in specification (2) with country-month fixed effect is positive. A 1%

increase in global institutional ownership is associated with an increase in total risk premium by 3.6 bps. This suggests that in DMs, global institutional ownership could actually increase firms' cost of capital by increasing their exposure to global risk. In addition, the institutional local risk premium decreases significantly with country level institutional ownership as well as the capital ratio. A 1% increase in $CountryIO$ is associated with a decrease of 3.4 bps in the institutional local risk premium in DMs and a 1% increase in CR is associated with a decrease of 48 bps in the institutional local risk premium. This is consistent with our model prediction that the institutional local risk premium declines with the risk-bearing capacity of financial institutions, either across countries proxied by $CountryIO$ or over time proxied by CR. We could also compare the coefficient of each risk premium on CR to see to what extent the risk premium depends on the financial conditions of global institutional investors. The attainable world premium is more negatively affected by CR. A 1% increase in CR is associated with a decrease of 1.72% in the attainable world market premium. In contrast, the local retail premium is less affected by the intermediary capital ratio and the effect of the same increase is only 37.7 bps. This reflects that the attainable world market premium and the institutional local risk premium depend on the risk-bearing capacity of institutional investors, whereas the retail local premium depends on the risk-bearing capacity of local retail investors. In summary, in DMs, we find that institutional ownership increases cost of capital in the cross-section.

Table VI reports the results of the same regressions for EMs. There is a stark contrast to the results in DMs. As in DMs, higher institutional ownership significantly predicts higher attainable world market risk premium. A 1% increase in firm-level IO predicts an increase in the attainable world market premium by 6.3 bps. Both the institutional local risk premium and the retail local risk premium are negatively and significantly predicted by firm-level IO. In specification (6) with country-time fixed effect, a 1% increase in IO is associated with a decrease in the institutional local risk premium by 4.3 bps. This result might be puzzling at the first sight because one would expect a stock to be more exposed to the institutional local factor if it has higher institutional ownership. But the institutional local factor is constructed as the return difference between the attainable market portfolio and its substitute portfolio of foreign institutional securities. The exposure of a stock to the institutional local factor could decrease with institutional ownership if higher institutional ownership increases its correlation with the substitute portfolio. As for the retail local risk premium, specification (8) with country-month fixed effect shows that a 1% increase in IO predicts a reduction of 10.2 bps in the retail local risk premium. This effect is economically important, which is consistent with our theory prediction that stocks with higher institutional ownership are more attainable and less exposed to the retail local factor, hence earn lower retail local risk premium due to better risk-sharing. One could argue that this is due to size effect because institutional ownership is higher in large stocks and large stocks tend to have lower risk premium.[27] It is worth noting that we discover this negative relationship between the retail local premium and IO after controlling for firm size, therefore we are not simply capturing the size effect. The net effect of a 1% increase in firm-level IO on total risk premium is a reduction of 8.1bps. This effect is statistically significant and economically important. At the country level, an increase of 1% in $CountryIO$ in EMs predict a decrease in the institutional local risk premium of 9.9bps, which is consistent with our theory that higher global institutional capacity in a country lowers the institutional local premium. Lastly, in EMs, only the attainable world market risk premium significantly and negatively predicted by CR. In summary, we find strong evidence that EM firms could reduce their cost of capital by increasing global institutional investment and this effect could not be explained by firm size and country-time fixed effects.

[27]It is well-documented that institutional investors have strong preference for large and visible stocks, for international evidence, see Ferreira and Matos (2008).

V. Concluding Remarks

We study how global institutional investment affects risk-sharing and risk premia in international equity markets. In order to do this, we develop and estimate an asset pricing model with home-biased retail investors and mandate-constrained global institutional investors. International risk-sharing depends on the coverage of institutional investors' mandate, their risk-bearing capacity and substitutability between institutional securities from different countries. In addition to an attainable world market premium, securities earn an institutional local risk premium due to imperfect risk-sharing across country borders, which decreases with the risk-bearing capacity of institutional investors. Retail securities earn a retail local premium due to imperfect risk-sharing between institutional and retail investors in each country. Existing international asset pricing theories focus on security-level investability and could not explain why investable securities are priced by local factors. We instead focus on the heterogeneous investment scopes of different investors. This allows us to provide a unique decomposition of market-level local risk premium into an institutional component and a retail component. Our theory provides a framework for analyzing the effect of global institutional investment on equity risk premium.

We estimate institutional and retail local premia using individual stock returns from 38 markets. First, the institutional local premium and the retail local premium are statistically significant and economically important in a wide range of DMs and EMs. The institutional local risk premium is lower in countries with higher global institutional ownership. Second, local risk premia are more important drivers of time-varying risk premium in EMs than in DMs. Third, higher firm-level global institutional ownership improves international risk-sharing and reduces the cost of capital in EMs.

Appendices

A. Motivation for model setup

This section contains empirical evidence in support of the model setup with home-biased retail investors and mandate-constrained institutional investors. Figure 6 shows the proportion of household equity and fund investment in home securities by the end of 2020 for 13 countries that have aggregate data on both household total and foreign equity investment. On average, households invest more than 80% of their portfolios at home.

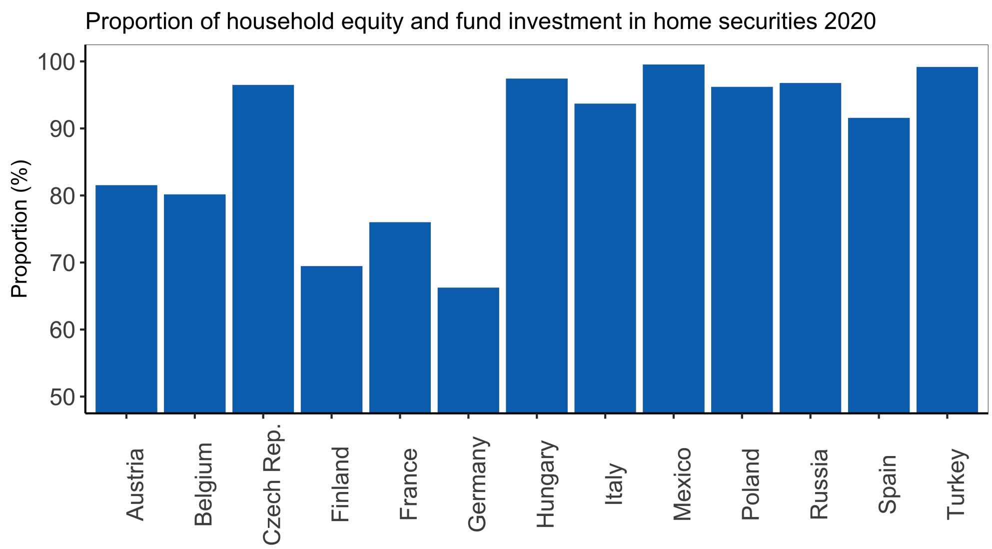

Figure 6. Proportion of household equity and fund investment in home securities 2020
This figure shows the proportion of househould equity and fund investment in home securities in 2020. It is calculated as one minus foreign investment divided by total investment. Househould total equity investment is sourced from OECD national account item 'Equity and investment fund shares/units' and household foreign investment is the 'Equity and investment fund shares' to the rest of the world from the IMF Coordinated Portfolio Investment Survey.

Figure 7 shows the proportion of all listed stocks included in FTSE all world indices by the end of 2018. FTSE indices include less than 20% of all the stocks in most countries.

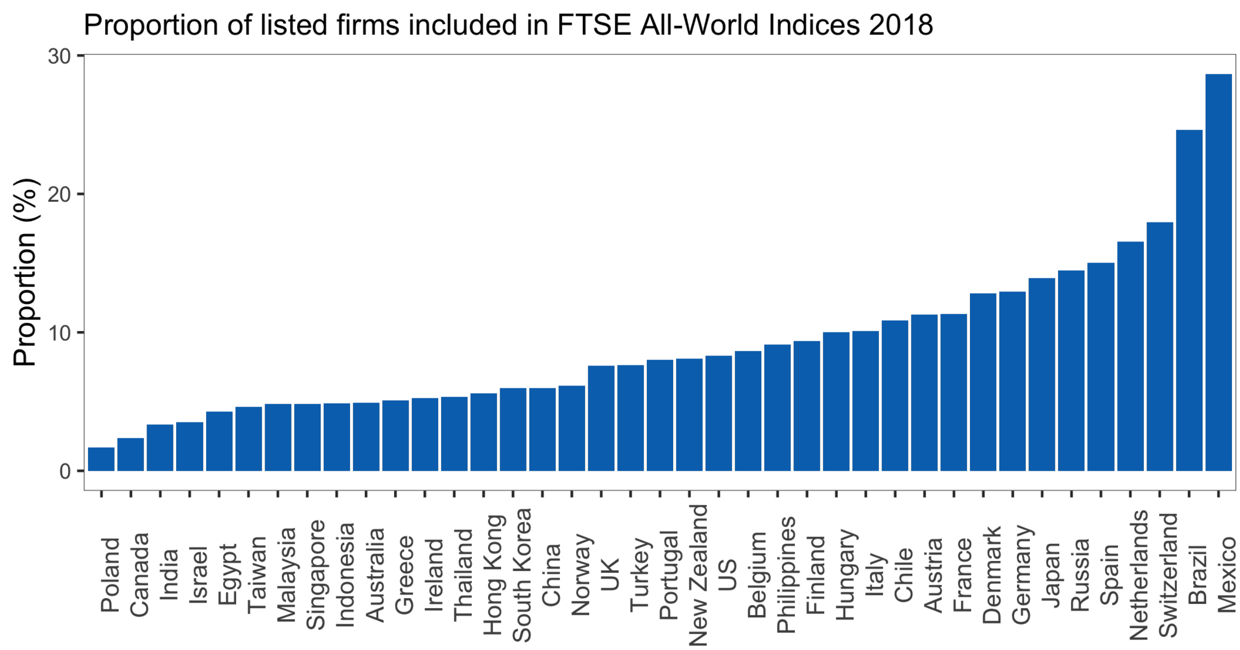

Figure 7. Proportion of listed stocks included in FTSE All-World indices in 2018
This figure plots the proportion of listed firms included in FTSE All-World indices. The number of included firms is from FTSE. The number of total listed firms is calculated as the number of listed firms that have a valid price in Compustat Global by the end of 2018.

B. Proof of model equilibrium

This appendix provides proof of the model equilibrium. Because the model is symmetric between the domestic and foreign countries, we focus on interpreting the results based on domestic securities and investors and introduce results about foreign securities and investors wherever necessary.

1. Problem setup

To simplify notation, we omit time subscripts unless necessary. Domestic and foreign securities are partitioned into four segments and we stack the excess returns of securities into a $N \times 1$ vector $\mathbf{r} = [\mathbf{r}'_R, \mathbf{r}'_I, \mathbf{r}'_{I^*}, \mathbf{r}'_{R^*}]'$. The market capitalization of risky securities are exogenous and we use $\mathbf{s}_R$, $\mathbf{s}_I$, $\mathbf{s}_{I^*}$ and $\mathbf{s}_{R^*}$ to denote a vector of the total supply of risky securities in terms of their market capitalization. For example, $\mathbf{s}_I = [s_{I_1}, \cdots, s_{I_{N_I}}]'$ is a $N_I \times 1$ vector of the market capitalization of domestic institutional securities.

Investor $k \in \{d, f, i\}$ has CARA utility and solves one-period mean-variance portfolio choice problem by choosing the optimal dollar investment $\mathbf{x}^k$ of securities in her choice set $\mathcal{C}^k$. Specifically $\mathcal{C}^d = \{R, I\}$, $\mathcal{C}^f = \{R^*, I^*\}$ and $\mathcal{C}^i = \{I, I^*\}$. The optimization problem investor k solves is:

$$\max_{x^k_{X\in\mathcal{C}^k}} \mathbb{E}[U(W^k_{t+1})] = \max_{x^k_{X\in\mathcal{C}^k}} \mathbb{E}[-\exp(-\gamma^k W^k_{t+1})] \tag{B.1}$$
$$W^k_{t+1} = W^k_t(1 + r_f) + \sum_{X\in\mathcal{C}^k} \mathbf{x}^{k'}_X \boldsymbol{r}_X$$

where γ^k is investor k's absolute risk aversion. X represents a market segment in investor k's choice set, for example for domestic retail investor $X \in \{R, I\}$. $\mathbf{x}^k_X = [x_1, \cdots, x_{N_X}]'$ is a $N_X \times 1$ vector containing investor k's dollar investment in securities in segment X, $\mathbf{r}_X = [r_{X_1}, \cdots, r_{X_{N_X}}]$ is a vector of one-period excess returns of assets in segment X, $\boldsymbol{\iota}$ is a vector of ones with the appropriate length.

With normally distributed return, the expected utility of investor k is:

$$\mathbb{E}[U(W^k_{t+1})] = -\exp\Big[-\gamma^k W^k_t(1 + r_f)\Big] \exp\Big[-\gamma^k \sum_{X\in\mathcal{C}^k} \mathbf{x}^{k'}_X \boldsymbol{\mu}_X + \frac{(\gamma^k)^2}{2} \sum_{X,Y\in\mathcal{C}^k} \mathbf{x}^{k'}_X V_{XY} \mathbf{x}^k_Y\Big] \tag{B.2}$$

where $\boldsymbol{\mu}_X$ is a $n_X \times 1$ vector of risk premia of securities in segment X. The optimization problem (B.1) is equivalent to:

$$\max_{x^k_{X\in\mathcal{C}^k}} \sum_{X\in\mathcal{C}^k} \mathbf{x}^{k'}_X \boldsymbol{\mu}_X - \frac{\gamma^k}{2} \sum_{X,Y\in\mathcal{C}^k} \mathbf{x}^{k'}_X V_{XY} \mathbf{x}^k_Y \tag{B.3}$$

The first order condition (FOC) of each investor's portfolio choice depends on their choice set $\mathcal{C}$s.

The FOC of the institutional investor i is:

$$V_{II}\mathbf{x}_I^i + V_{II^*}\mathbf{x}_{I^*}^i = \frac{1}{\gamma^i}\boldsymbol{\mu}_I \tag{B.4}$$

$$V_{I^*I}\mathbf{x}_I^i + V_{I^*I^*}\mathbf{x}_{I^*}^i = \frac{1}{\gamma^i}\boldsymbol{\mu}_{I^*} \tag{B.5}$$

The FOC of the domestic retail investor d is:

$$V_{RR}\mathbf{s}_R + V_{RI}\mathbf{x}_I^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_R \tag{B.6}$$

$$V_{IR}\mathbf{s}_R + V_{II}\mathbf{x}_I^d = \frac{1}{\gamma^d}\boldsymbol{\mu}_I \tag{B.7}$$

The FOC of the foreign retail investor f:

$$V_{I^*I^*}\mathbf{x}_{I^*}^f + V_{I^*R^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} \tag{B.8}$$

$$V_{R^*I^*}\mathbf{x}_{I^*}^f + V_{R^*R^*}\mathbf{s}_{R^*} = \frac{1}{\gamma^f}\boldsymbol{\mu}_{R^*} \tag{B.9}$$

We used in the FOCs of domestic and foreign retail investors the following market-clearing conditions for domestic and foreign retail securities:

$$\mathbf{x}_R^d = \mathbf{s}_R \tag{B.10}$$

$$\mathbf{x}_{R^*}^f = \mathbf{s}_{R^*} \tag{B.11}$$

From (B.7) and (B.8), we could express the dollar demand for institutional securities by retail investors in terms of the risk premia of institutional securities.

$$\mathbf{x}_I^d = V_{II}^{-1}[\frac{1}{\gamma^d}\boldsymbol{\mu}_I - V_{IR}\mathbf{s}_R] \tag{B.12}$$

$$\mathbf{x}_{I^*}^f = V_{I^*I^*}^{-1}[\frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} - V_{I^*R^*}\mathbf{s}_{R^*}] \tag{B.13}$$

Substituting in (B.12) into (B.6) yields an expression of the risk premia of domestic retail securities in terms of the risk premia of domestic institutional securities:

$$\boldsymbol{\mu}_R = V_{RI} V_{II}^{-1} \boldsymbol{\mu}_I + \gamma^d (V_{RR} - V_{RI} V_{II}^{-1} V_{IR}) \mathbf{s}_R \tag{B.14}$$

The risk premia of domestic retail securities have two components. The first term $V_{RI} V_{II}^{-1} \boldsymbol{\mu}_I$ reflects the risk premia earned by retail securities for their returns spanned by their institutional counterpart. Intuitively, the pricing of retail securities is "benchmarked" against their institutional counterpart and the level of this risk premium depends on how well domestic retail securities can be spanned by domestic institutional securities. $V_{II}^{-1} V_{IR}$ is the regression coefficient of a multiple regression of domestic retail securities on domestic institutional securities. $V_{RR} - V_{RI} V_{II}^{-1} V_{IR}$ measures the residual covariance among domestic retail securities that cannot be explained by this linear regression. The second term is compensation for this residual covariance that depends on the absolute risk aversion of the retail investor γ^d.

To get equilibrium risk premium and portfolio holdings we invoke the market-clearing conditions of domestic and foreign institutional securities:

$$\mathbf{x}_I^d + \mathbf{x}_I^i = \mathbf{s}_I \tag{B.15}$$

$$\mathbf{x}_{I^*}^f + \mathbf{x}_{I^*}^i = \mathbf{s}_{I^*} \tag{B.16}$$

2. Equilibrium risk premia

Using the market clearing conditions of institutional securities (B.15, B.16), we get the dollar investment in institutional securities by the institutional investor as the residual of market supply after subtracting the demand from retail investors:

$$\mathbf{x}_I^i = \mathbf{s}_I - V_{II}^{-1}[\frac{1}{\gamma^d}\boldsymbol{\mu}_I - V_{IR}\mathbf{s}_R] \tag{B.17}$$

$$\mathbf{x}_I^i = \mathbf{s}_{I^*} - V_{I^*I^*}^{-1}[\frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} - V_{I^*R^*}\mathbf{s}_{R^*}] \tag{B.18}$$

Substituting institutional investor's investment in institutional securities (B.17) and (B.18) into its FOC (B.4) and (B.5) gives us a linear system from which we solve for institutional risk premia $\boldsymbol{\mu}_I$ and $\boldsymbol{\mu}_{I^*}$:

$$\frac{1}{\gamma^i}\boldsymbol{\mu}_I = V_{II}\left[\mathbf{s}_I - V_{II}^{-1}(\frac{1}{\gamma^d}\boldsymbol{\mu}_I - V_{IR}\mathbf{s}_R)\right] + V_{II^*}\left[\mathbf{s}_{I^*} - V_{I^*I^*}^{-1}(\frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} - V_{I^*R^*}\mathbf{s}_{R^*})\right] \tag{B.19}$$

$$\frac{1}{\gamma^i}\boldsymbol{\mu}_{I^*} = V_{I^*I}\left[\mathbf{s}_I - V_{II}^{-1}(\frac{1}{\gamma^d}\boldsymbol{\mu}_I - V_{IR}\mathbf{s}_R)\right] + V_{I^*I^*}\left[\mathbf{s}_{I^*} - V_{I^*I^*}^{-1}(\frac{1}{\gamma^f}\boldsymbol{\mu}_{I^*} - V_{I^*R^*}\mathbf{s}_{R^*})\right] \tag{B.20}$$

We rewrite the linear system (B.19) and (B.20)more compactly in matrix form:

$$\begin{bmatrix} \frac{1}{\gamma^D}\mathbf{I} & \frac{1}{\gamma^f}V_{II^*}V_{I^*I^*}^{-1} \\ \frac{1}{\gamma^d}V_{I^*I}V_{II}^{-1} & \frac{1}{\gamma^F}\mathbf{I} \end{bmatrix} \begin{bmatrix} \boldsymbol{\mu}_I \\ \boldsymbol{\mu}_{I^*} \end{bmatrix} = \begin{bmatrix} V_{II}\mathbf{s}_I + V_{IR}\mathbf{s}_R + V_{II^*}\mathbf{s}_{I^*} + V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*} \\ V_{I^*I}\mathbf{s}_I + V_{I^*R^*}\mathbf{s}_{R^*} + V_{I^*I^*}\mathbf{s}_{I^*} + V_{I^*I}V_{II}^{-1}V_{IR}\mathbf{s}_R \end{bmatrix} \tag{B.21}$$

where we define the average risk aversion of investors in the domestic and foreign markets γ^D and γ^F such that:

$$\frac{1}{\gamma^D} = \frac{1}{\gamma^d} + \frac{1}{\gamma^i} \tag{B.22}$$

$$\frac{1}{\gamma^F} = \frac{1}{\gamma^f} + \frac{1}{\gamma^i} \tag{B.23}$$

Rearranging the linear system (B.21), we get an expression of the risk premia of institutional securities in terms of exogenous inputs:

$$\begin{bmatrix} \boldsymbol{\mu}_I \\ \boldsymbol{\mu}_{I^*} \end{bmatrix} = \begin{bmatrix} V_{II} & \mathbf{O} \\ \mathbf{O} & V_{I^*I^*} \end{bmatrix} \begin{bmatrix} \frac{1}{\gamma^D} V_{II} & \frac{1}{\gamma^f} V_{II^*} \\ \frac{1}{\gamma^d} V_{I^*I} & \frac{1}{\gamma^F} V_{I^*I^*} \end{bmatrix}^{-1} \begin{bmatrix} V_{II}\mathbf{s}_I + V_{IR}\mathbf{s}_R + V_{II^*}\mathbf{s}_{I^*} + V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*} \\ V_{I^*I}\mathbf{s}_I + V_{I^*R^*}\mathbf{s}_{R^*} + V_{I^*I^*}\mathbf{s}_{I^*} + V_{I^*I}V_{II}^{-1}V_{IR}\mathbf{s}_R \end{bmatrix} \tag{B.24}$$

$$= \begin{bmatrix} V_{II} & \mathbf{O} \\ \mathbf{O} & V_{I^*I^*} \end{bmatrix} \begin{bmatrix} \gamma^D \Phi_I^{-1} & -\frac{\gamma^D\gamma^F}{\gamma^f} \Phi_I^{-1} V_{II^*} V_{I^*I^*}^{-1} \\ -\frac{\gamma^D\gamma^F}{\gamma^d} \Phi_{I^*}^{-1} V_{I^*I} V_{II}^{-1} & \gamma^F \Phi_{I^*}^{-1} \end{bmatrix} \begin{bmatrix} V_{II}\mathbf{s}_I + V_{IR}\mathbf{s}_R + V_{II^*}\mathbf{s}_{I^*} + V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*} \\ V_{I^*I}\mathbf{s}_I + V_{I^*R^*}\mathbf{s}_{R^*} + V_{I^*I^*}\mathbf{s}_{I^*} + V_{I^*I}V_{II}^{-1}V_{IR}\mathbf{s}_R \end{bmatrix}$$

where Φ_I and Φ_{I^*} are defined as follows:

$$\Phi_I = V_{II} - \frac{\gamma^D\gamma^F}{\gamma^d\gamma^f} V_{II^*} V_{I^*I^*}^{-1} V_{I^*I} \tag{B.25}$$

$$\Phi_{I^*} = V_{I^*I^*} - \frac{\gamma^D\gamma^F}{\gamma^d\gamma^f} V_{I^*I} V_{II}^{-1} V_{II^*} \tag{B.26}$$

Φ_I is the conditional covariance among domestic institutional securities that is not explained by the span of foreign institutional securities. Unlike the conditional covariance matrices in (B.14), here the amount of conditioning is adjusted by a factor of $\frac{\gamma^D\gamma^F}{\gamma^d\gamma^f}$. The higher this ratio, the lower the risk-bearing capacity of the institutional investor relative to retail investors.[28]

To simplify the conditional covariance matrix Φ_I (B.25), we assume that the following approximation holds:

$$\Phi_I^{-1} \approx \theta V_{II}^{-1} \tag{B.27}$$

where $\theta > 1$ is a scalar that increases if domestic and foreign institutional securities could better substitute each other.

Approximation (B.27) holds exactly when there is only one institutional and one retail

[28] To see this $\frac{\gamma^D\gamma^F}{\gamma^d\gamma^f} = \frac{(\gamma^i)^2}{(\gamma^d+\gamma^i)(\gamma^f+\gamma^i)}$, which increases with the risk aversion of the institutional investor γ^i hence decreases with the risk-bearing capacity of the institutional investor.

security in each country. To interpret θ, suppose also $corr(r_I, r_{I^*}) = \rho$, then:

$$V_{II^*}V_{I^*I^*}^{-1}V_{I^*I} = \rho^2 V_{II}$$

$\rho^2 < 1$ captures the fact that domestic and foreign institutional securities are not perfect substitutes. It measures the limited substitutability between securities in segmented markets. When $\rho^2 = 1$, $V_{II^*}V_{I^*I^*}^{-1}V_{I^*I} = V_{II}$ and $V_{I^*I}V_{II}^{-1}V_{II^*} = V_{I^*I^*}$, domestic and foreign institutional securities perfectly span each other. In this case, θ can be expressed as:

$$\theta = [1 - \frac{\gamma^D\gamma^F}{\gamma^d\gamma^f}\rho^2]^{-1} \tag{B.28}$$

θ is influenced by the institutional investor's decision to trade off risk across domestic and foreign institutional securities. This is determined by both ρ^2 as well as $\frac{\gamma^D\gamma^F}{\gamma^d\gamma^f}$. ρ^2 measures how much diversification the institutional investor enjoys by investing in both domestic and foreign institutional securities. When the absolute correlation is higher, domestic and foreign institutional securities better substitute each other, the institution has less diversification and the higher θ is. When retail investors's relative wealth with respect to the institutional investor increases, $\frac{\gamma^D\gamma^F}{\gamma^d\gamma^f}$ increases and θ increases.

Solving for the risk premia of domestic institutional securities, using the approximation

(B.27) , we have:

$$\boldsymbol{\mu}_I = \gamma^D V_{II}\Phi_I^{-1}\Big[(V_{II} - \frac{\gamma^F}{\gamma^f}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I})\mathbf{s}_I + (V_{IR} - \frac{\gamma^F}{\gamma^f}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR})\mathbf{s}_R + (1-\frac{\gamma^F}{\gamma^f})V_{II^*}\mathbf{s}_{I^*} \tag{B.29}$$

$$+ (1-\frac{\gamma^F}{\gamma^f})V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*}\Big]$$

$$= \gamma\big[V_{II}\mathbf{s}_I + V_{IR}\mathbf{s}_R + V_{II^*}\mathbf{s}_{I^*} + V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*}\big]$$

$$+ \frac{\gamma^i}{\gamma^f}\gamma\big[(V_{II} - V_{II^*}V_{I^*I^*}^{-1}V_{I^*I})\mathbf{s}_I + (V_{IR} - V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR})\mathbf{s}_R\big]$$

Substituting (B.29) into (B.14) gives us the risk premium of domestic retail securities:

$$\boldsymbol{\mu}_R = \gamma V_{RI}V_{II}^{-1}\big[V_{II}\mathbf{s}_I + V_{IR}\mathbf{s}_R + V_{II^*}\mathbf{s}_{I^*} + V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*}\big] \tag{B.30}$$

$$+ \frac{\gamma^i}{\gamma^f}\gamma V_{RI}V_{II}^{-1}\big[(V_{II} - V_{II^*}V_{I^*I^*}^{-1}V_{I^*I})\mathbf{s}_I + (V_{IR} - V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR})\mathbf{s}_R\big] + \gamma^d(V_{RR} - V_{II}^{-1}V_{IR})\mathbf{s}_R$$

where we use the fact that $1 - \frac{\gamma^F}{\gamma^f} = \frac{\gamma^F}{\gamma^i}$.

γ is the aggregate risk aversion of the economy:

$$\gamma = \frac{\gamma^D\gamma^F}{\gamma^i}\theta \tag{B.31}$$

γ is easier to interpret in the special case when there is only one institutional security in each country. Assume that the correlation between domestic and foreign institutional securities is $corr(r_I, r_{I^*}) = \rho$, the aggregate risk aversion γ is:

$$\frac{1}{\gamma} = \frac{\gamma^i}{\gamma^D\gamma^F}\theta^{-1} \tag{B.32}$$

$$= \frac{(\gamma^d+\gamma^i)(\gamma^f+\gamma^i)}{\gamma^d\gamma^f\gamma^i}\frac{\gamma^d\gamma^f + \gamma^d\gamma^i + \gamma^f\gamma^i + (1-\rho^2)(\gamma^i)^2}{(\gamma^d+\gamma^i)(\gamma^f+\gamma^i)}$$

$$= \frac{1}{\gamma^d} + \frac{1}{\gamma^f} + \frac{1}{\gamma^i} + (1-\rho^2)\frac{\gamma^i}{\gamma^d\gamma^f}$$

When $\rho^2 = 1$, the aggregate risk-aversion is the average risk aversion of the three representative investors in a frictionless economy. When $\rho^2 \neq 1$, the aggregate risk-aversion decreases as ρ^2 decreases. Intuitively, in this economy, the aggregate risk-aversion depends on the correlation between domestic and foreign institutional securities. As ρ^2 decreases, institutional investors enjoy more international diversification and the aggregate risk aversion is lower.

We define the *attainable return* of any domestic security j as the fitted value of regressing its return r_j onto domestic institutional securities:

$$\hat{r}_j = B_{jI}\mathbf{r}_I \tag{B.33}$$

where $B_{jI} = V_{jI}V_{II}^{-1}$ is a $1 \times N_I$ row vector of the multiple regression coefficient of regressing the return of security j onto domestic institutional securities.

We define the *attainable domestic market portfolio* $\hat{D}$ as the portfolio of domestic institutional securities that optimally mimics the domestic market portfolio. The return on this portfolio is the fitted value of a multiple linear regression of the domestic market portfolio on domestic institutional securities:

$$r_{\hat{D}} = \boldsymbol{\omega}_I'\mathbf{r}_I + \boldsymbol{\omega}_R' B_{RI}\mathbf{r}_I \tag{B.34}$$

.

where $B_{RI} = V_{RI}V_{II}^{-1}$ is a $n_R \times n_I$ matrix of multiple regression coefficients of retail securities on institutional securities and $\boldsymbol{\omega}_I$ and $\boldsymbol{\omega}_R$ are vectors of the weight of domestic institutional and retail securities in the domestic market portfolio. The *attainable foreign market portfolio* is defined in the same way. We define the value-weighted portfolio of the attainable domestic portfolio ($\hat{D}$) and the attainable foreign portfolio ($\hat{F}$) as the *attainable* *world market portfolio* ($\hat{W}$):

$$r_{\hat{W}} = \omega_D r_{\hat{D}} + \omega_F r_{\hat{F}} \tag{B.35}$$

where $\omega_D = \frac{M_D}{M_W}$ and $\omega_F = \frac{M_F}{M_W}$ are the weight of domestic and foreign market portfolios in the world market portfolio. The attainable world market portfolio represents the world market exposure that is attainable by investing in domestic and foreign institutional securities.

We define the *substitute portfolio* for the attainable domestic market portfolio as the portfolio of foreign institutional securities that optimally mimics the attainable domestic market portfolio. The return on this portfolio is the fitted value of a multiple linear regression of the attainable domestic market portfolio onto foreign institutional securities:

$$r_{D^s} = \boldsymbol{\omega}_I' B_{II^*} \mathbf{r}_{I^*} + \boldsymbol{\omega}_R' B_{RI} B_{II^*} \mathbf{r}_{I^*} \tag{B.36}$$

where $B_{II^*} = V_{II^*} V_{I^*I^*}^{-1}$ is a $n_I \times n_{I^*}$ matrix of multiple regression coefficient of domestic institutional securities on foreign institutional securities.

From (B.29), the equilibrium risk premia of any domestic institutional security I_j can be expressed in terms of their covariance with aggregate factor portfolios as follows:

$$\begin{aligned}
\mu_{I_j} &= V_{I_jI}\mathbf{s}_I + V_{I_jR}\mathbf{s}_R + V_{I_jI^*}\mathbf{s}_{I^*} + V_{I_jI^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*} \\
&\quad + \frac{\gamma^i}{\gamma^f}\gamma\left[(V_{I_jI} - V_{I_jI^*}V_{I^*I^*}^{-1}V_{I^*I})\mathbf{s}_I + (V_{I_jR} - V_{I_jI^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR})\mathbf{s}_R\right] \\
&= M_D cov(r_{I_j}, \boldsymbol{\omega}_I'\mathbf{r}_I + \boldsymbol{\omega}_R'\mathbf{r}_R) + M_F cov(r_{I_j}, \boldsymbol{\omega}_{I^*}'\mathbf{r}_{I^*} + \boldsymbol{\omega}_{R^*}' B_{R^*I^*}\mathbf{r}_{I^*}) \\
&\quad + \frac{\gamma^i}{\gamma^f}\gamma M_D cov(r_{I_j}, \boldsymbol{\omega}_I'\mathbf{r}_I + \boldsymbol{\omega}_R'\mathbf{r}_R - \boldsymbol{\omega}_I' B_{II^*}\mathbf{r}_{I^*} - \boldsymbol{\omega}_R' B_{RI} B_{II^*}\mathbf{r}_{I^*}) \\
&= \gamma M_W cov(r_{I_j}, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(r_{I_j}, r_{\hat{D}} - r_{D^s}\big)
\end{aligned} \tag{B.37}$$

where we use the fact that $cov(r_I, r_D) = cov(r_I, r_{\hat{D}})$. Domestic institutional securities earn two risk premia. The first one is an attainable world risk premium for their covariance with the attainable world market return $r_{\hat{W}}$, which we also refer to as the *attainable world market factor* $f^{\hat{W}}$. The second one is for their covariance with the return difference between the attainable domestic market portfolio and its substitute portfolio of foreign securities, which we define as the *institutional local factor*:

$$f^{ilocal} = r_{\hat{D}} - r_{D^s} \tag{B.38}$$

From (B.30), the equilibrium risk premia of any domestic retail security R_j can be expressed similarly in terms of its covariances with factor portfolios:

$$\mu_{R_j} = \gamma M_W cov(\hat{r}_{R_j}, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(\hat{r}_{R_j}, r_{\hat{D}} - r_{D^s}\big) \tag{B.39}$$
$$+ \gamma^d M_R cov\big(r_{R_j}, r_R - \hat{r}_R\big)$$

where $\hat{r}_{R_j}$ is the attainable return of security R_j defined in (B.33).

Domestic retail securities earn a third risk premium for their covariance with the residual retail risk that is orthogonal to the span of institutional securities, which we define as the *retail local factor*:

$$f^{rlocal} = r_R - \hat{r}_R \tag{B.40}$$

The risk premium of any domestic security j can be expressed as follows:

$$\mu_j = \gamma M_W cov(\hat{r}_j, r_{\hat{W}}) + \frac{\gamma^i}{\gamma^f}\gamma M_D cov\big(\hat{r}_j, r_{\hat{D}} - r_{D^s}\big) + \gamma^d M_R cov\big(r_j - \hat{r}_j, r_R - \hat{r}_R\big) \tag{B.41}$$

where we used the fact that for institutional securities $r_{I_j} = \hat{r}_{I_j}$ and for retail securities $cov(\hat{r}_{R_j}, r_R - \hat{r}_R) = 0$.

The price of security j's attainable world market risk γM_W is the same across four segments. The price of security j's institutional local risk is $\frac{\gamma^i}{\gamma^f}\gamma M_D$, which increases when the institutional investor becomes more risk averse relative to the foreign retail investor. The intuition is that when the institutional investor becomes more risk averse relative to the foreign retail investor ($\gamma^f \downarrow, \gamma^i \uparrow$), foreign securities are held more by the foreign retail investor and less by the institution. This change in ownership composition results in less international risk sharing hence higher institutional local premium relative to the attainable world premium. γ^i also has a direct level effect on the overall level of risk premium through the aggregate risk aversion γ, as can be seen from (B.32).

Consider the special case in which there is only one institutional security and one retail security in each country. Suppose the correlation between domestic and foreign institutional securities is $corr(r_I, r_{I^*}) = \rho$, the correlation between domestic retail and institutional security is $corr(r_R, r_I) = \rho_R$ and the correlation between foreign retail and institutional security is $corr(r_{R^*}, r_{I^*}) = \rho_R^*$. For any domestic security j, suppose its volatility is σ_j and its correlation with the domestic institutional security is $corr(r_j, r_I) = \rho_j$. Then its attainable return is $\hat{r}_j = \rho_j \frac{\sigma_j}{\sigma_I} r_I$. Simplifying (B.41), the risk premium of j can be expressed as:

$$\mu_j = \gamma M_W \rho_j \frac{\sigma_j}{\sigma_I} cov(r_I, r_{\hat{W}}) + \gamma \frac{\gamma^i}{\gamma^f} M_D \rho_j \frac{\sigma_j}{\sigma_I}(1-\rho^2) cov(r_I, r_{\hat{D}}) + \gamma^d M_R (1-\rho_j^2)\sigma_j \sigma_R \quad \text{(B.42)}$$

In this simplified case, domestic security j earns attainable world and institutional local risk premium through their correlation with the domestic institutional security ρ_j. The novelty in our model is the indirect covariance between domestic and foreign retail securities embedded in the attainable world risk premium of the domestic retail security: $cov(\hat{r}_R, \hat{r}_{R^*}) = \rho_R \rho \rho_{R^*} \sigma_R \sigma_{R^*}$. These two remote markets do not share a single common investor yet share risk through a chain of correlations: the correlation between domestic retail and institutional securities ρ_R, the correlation between domestic and foreign institutional securities ρ and the correlation between foreign institutional and retail securities ρ_{R^*}. The relative importance of global versus local risk premiums depends on three Cs. First, the *Correlation* between domestic and foreign institutional securities ρ, which determines the quantity of risk that can be shared internationally by trading institutional securities; second, the risk-bearing *Capacity* of the institutional investor, $1/\gamma^i$, which determines the level of the institutional local risk premium; third, the *Coverage* of the institution's choice set represented by ρ_j. The correlation between a domestic security and the domestic institutional security reflects to what extent its exposure can be spanned by the institutional security (covered by the institution's mandate). Since the retail local risk premium is compensation for the unspanned retail risk, the higher the correlation, the lower the retail local risk premium. $\rho_j = 1$ when either j is an institutional security or because j is a retail security that is perfectly correlated with the domestic institutional security. In this case, the retail local risk premium becomes zero. Notably, the correlation between a domestic retail security and foreign securities do not enter equilibrium risk premium because international risk-sharing in the model is channeled through its correlation with the domestic institutional security.

Beta representation

Apply B.41 to the attainable world market factor, the institutional local factor and the retail local factor. Because these three factors are constructed to be orthogonal to each other, we have:

$$\mu^{\hat{W}} = \gamma M_W var(f^{\hat{W}}) \tag{B.43}$$

$$\mu^{ilocal} = \frac{\gamma^i}{\gamma^f}\gamma M_D var(f^{ilocal}) \tag{B.44}$$

$$\mu^{rlocal} = \gamma^d M_R var(f^{rlocal}) \tag{B.45}$$

Substitute (B.43), (B.44) and (B.45) back into (B.41) yields the following beta representation of the risk premium of domestic security j:

$$\mu_j = \beta^{\hat{W}}\mu^{\hat{W}} + \beta_j^{ilocal}\mu^{ilocal} + \beta_j^{rlocal}\mu^{rlocal} \tag{B.46}$$

where

$$\beta_j^{\hat{W}} = \frac{cov(\hat{r}_j, f^{\hat{W}})}{var(f^{\hat{W}})} \tag{B.47}$$

$$\beta_j^{ilocal} = \frac{cov(\hat{r}_j, f^{ilocal})}{var(f^{ilocal})} \tag{B.48}$$

$$\beta_j^{rlocal} = \frac{cov(r_j - \hat{r}_j, f^{rlocal})}{var(f^{rlocal})} \tag{B.49}$$

3. Equilibrium investment

Substituting (B.29) into the FOC of the domestic retail investor (B.12), we get the holdings of domestic institutional securities by the domestic retail investor:

$$\mathbf{x}_I^d = \frac{\gamma^D}{\gamma^d}\Phi_I^{-1}\Bigg[(V_{II} - \frac{\gamma^F}{\gamma^f}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I})\mathbf{s}_I + (V_{IR} - \frac{\gamma^F}{\gamma^f}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR})\mathbf{s}_R + (1 - \frac{\gamma^F}{\gamma^f})V_{II^*}\mathbf{s}_{I^*}$$
$$+ (1 - \frac{\gamma^F}{\gamma^f})V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*}\Bigg] - V_{II}^{-1}V_{IR}\mathbf{s}_R \tag{B.50}$$

Using the approximation (B.27), we simplify the equilibrium investment of the domestic retail investor in domestic institutional securities (B.50):

$$\mathbf{x}_I^d = \frac{\gamma^D}{\gamma^d}\theta\underbrace{\left[\mathbf{s}_I + V_{II}^{-1}V_{IR}\mathbf{s}_R\right]}_{\hat{D}} + \frac{\gamma^D\gamma^F}{\gamma^d\gamma^i}\theta\underbrace{\left[V_{II}^{-1}V_{II^*}\mathbf{s}_{I^*} + V_{II}^{-1}V_{II^*}V_{I^*I^*}^{-1}V_{I^*R^*}\mathbf{s}_{R^*}\right]}_{F^s} - \underbrace{V_{II}^{-1}V_{IR}\mathbf{s}_R}_{\hat{R}}$$
$$- \frac{\gamma^D\gamma^F}{\gamma^d\gamma^f}\theta\underbrace{\left[V_{II}^{-1}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}\mathbf{s}_I + V_{II}^{-1}V_{II^*}V_{I^*I^*}^{-1}V_{I^*I}V_{II}^{-1}V_{IR}\mathbf{s}_R\right]}_{H} \tag{B.51}$$

where we use the fact that $1 - \frac{\gamma^F}{\gamma^f} = \frac{\gamma^F}{\gamma^i}$.

We define the following constants A^d, A^f, γ as:

$$A^d = \frac{\gamma^D}{\gamma^d}\theta \tag{B.52}$$

$$A^f = \frac{\gamma^F}{\gamma^f}\theta \tag{B.53}$$

where A^d (A^f) measures the relative risk-bearing capacity the domestic (foreign) retail investor with respect to the institutional investor.

By market clearing for the domestic retail securities, the domestic retail investor holds the domestic retail portfolio:

$$\mathbf{x}_R^d = \mathbf{s}_R \tag{B.54}$$

We could express domestic retail investor's dollar investment in domestic securities in terms of its four components:

$$\mathbf{x}_I^d + \mathbf{s}_R^d = \underbrace{A^d \hat{D}}_{\text{attainable domestic}} + \underbrace{D - \hat{D}}_{\text{unattainable local risk}} + \underbrace{\frac{\gamma}{\gamma^d}}_{\text{substitute for foreign}} F^s - \frac{\gamma^i}{\gamma^d \gamma^f}\gamma H \tag{B.55}$$

where H is a portfolio of domestic institutional securities that optimally mimics the substitute portfolio for domestic investment D^s. This position is a hedge that the domestic retail investor provides to the institutional investor for their tilt away from the substitute portfolio for domestic investment D^s in the foreign country.[29]

The size of this position depends on the level of correlation between domestic and foreign institutional securities. To see this, consider the special case in which there is only one institutional security and one retail security in each country, assuming that the correlation between the domestic institutional security and the foreign institutional security is $corr(r_I, r_{I^*}) = \rho$,

[29]Because the foreign retail investor tilts toward and the institutional investor tilts away from the foreign substitute portfolio, it is as if the institution "sells" the substitute portfolio to the foreign retail investor.

H can be simplified as:

$$H = \rho^2 \hat{D} \tag{B.56}$$

Unless the correlation between domestic and foreign institutional securities is very high, the magnitude of this term is very small compared to the other components of investors' portfolio.

Using the market clearing condition for domestic institutional securities, we get the dollar investment in domestic institutional securities by the institutional investor:

$$\mathbf{x}_I^i = (1 - A^d)\hat{D} - \frac{\gamma}{\gamma^d} F^s + \frac{\gamma^i}{\gamma^d \gamma^f} \gamma H \tag{B.57}$$

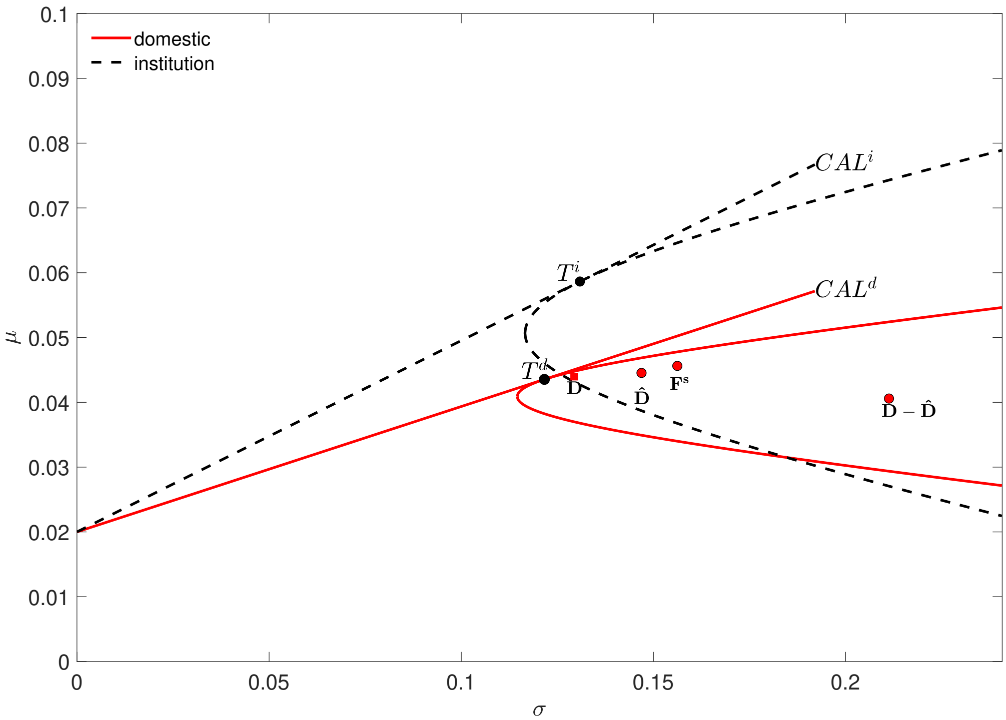

Figure 8. Efficient frontiers of the domestic retail investor and the institutional investor
This figure plots the efficient frontiers of the domestic retail investor and the institutional investor. T^i and T^d are the tangency portfolios of the institutional investor and the domestic retail investor. CAL^i and CAL^d are the capital allocation lines of the two investors.

Figure 8 shows the portfolio frontiers of the domestic retail investor and the institutional investor. Unlike existing partial segmentation models in which the frontier of a unrestricted investor is superior than the frontier of a restricted investor, the portfolio frontiers of two investors intersect.

4. The effect of global institutional investment on equilibrium risk premia

How does the risk premium of a domestic retail stock j change after it is included in the choice set of global institutions? On the one hand, it would enjoy direct international risk-sharing and no longer earn the retail local risk premium. On the other hand, the attainable world market risk premium and the institutional local risk premium now depends on the covariance between the raw return of the security (r_j) with the risk factors rather than its attainable component ($\hat{r}_j$). Taking the difference between the risk premium of an institutional security j (B.39) and a retail security j (B.37) yields the change in its risk premium:[30]

$$\Delta\mu_j \approx \gamma cov(r_j - \hat{r}_j, r_{\hat{F}})M_F - \gamma^d cov(r_j - \hat{r}_j, r_R)M_R \tag{B.58}$$

If a retail security can be perfectly replicated by its institutional counterpart, then inclusion in the institutional choice set has no effect on its risk premium. Otherwise, the effect of inclusion depends on how its non-attainable exposure covaries with retail local securities and with foreign attainable returns. Securities that covary more with retail local securities and less with foreign attainable returns get more reduction in risk premium post-inclusion.

[30] Here we assume that the institutional local risk factor lies in the span of domestic institutional securities.

C. Data construction

1. Stock universe

We start with the universe of securities from 48 countries included in the FTSE All World Index, and we exclude countries that have less than 10 stocks with positive global institutional ownership by the end of 2003. This leaves us with a total of 38 countries in our sample. [31] We apply the following filters for securities to be included in our sample following Chaieb, Langlois, and Scaillet (2021)

1. Securities that are ordinary shares or depository receipt (tpci='0', 'F')
2. We remove non-common stocks based on the keywords used in Griffin, Kelly, and Nardari (2010), Lee (2011) according to securities' issue description (*dsci*).
3. We only keep securities that are the major security of their company. In Compustat, a security is a major security at a given time if its security identifier (iid) matches the value of the major security item (for companies that are located in the US and Canada (loc="USA","CAN"), the primary security item is "PRIHISTUSA" and "PRIHISTCAN" and for companies from the rest of the world, the primary security item is "PRIHISTROW").

2. Security level variable calculation

1. **Total return index**: we calculate the total return index of securities as $prccd/ajexdi * trfd$, we set $trfd$ to 1 if it is missing and we use currency $curcdd$ and exchange rate from *exrt_dly* to convert total return indices to USD. We apply a delisting return of -30% when delisting is performance related (*dlrsni*). We remove erroneous returns that are identified as:
    - A return is set to missing if the absolute value of it is greater than 2

[31] We exclude Colombia, Czech Republic, Egypt, Hungary, Kuwait, Pakistan, Qatar, Romania, Saudi Arabia and United Arab Emirates.

- A return is set to missing if it is less than -1
- If the absolute value of a return is greater than 1 and the absolute value of one-period lagged return is greater than 1, and the absolute value of the geometric average return is less than 20%, the return is set to missing.
- To further limit the effect of outliers, we winsorize return observations at the 1% and 99% levels in each month for each country.

2. **Market capitalization**: we calculate market capitalization as $prccd * cshoc/qunit$ from *g_secd* for non-North American stocks. For North American stocks, we calculate market capitalization as $prccd * cshoi$. Last report number of shares outstanding is from *sec_afnd*. We convert market capitalization that is not denominated in USD into USD.
3. **Institutional ownership**: we follow the SAS code of Ferreira and Matos (2008) to calculate firm-level institutional ownership. We also calculate the country and region weight of each institution in each quarter. We follow Bartram et al. (2015) and define global institutions as institutions whose maximum weight invested in a country in a given quarter does not exceed 90% and maximum weight invested in a region in a given quarter does not exceed 80%.
4. **Book-to-market ratio**: following Davis, Fama, and French (2000), we calculate book equity as stockholders' book equity, plus balance sheet deferred taxes and investment tax credit (Compustat annual item TXDITC) if available, minus the book value of preferred stock. Stockholders' equity is the value reported by Compustat (item SEQ), if it is available. If not, we measure stockholders' equity as the book value of common equity (item CEQ) plus the par value of preferred stock (item PSTK), or the book value of assets (item AT) minus total liabilities (item LT). Depending on availability, we use redemption (item PSTKRV), liquidating (item PSTKL), or par value (item PSTK) for the book value of preferred stock. We calculate monthly book-to-market ratio using

the ratio between the last reported book value in December each year and the market capitalization of each month. We winsorize BM ratio at 1% and 99%.

5. **Dividend yield**: we calculate the annual dividend of each security as the sum of div in the security daily table in a year. We calculate the monthly dividend yield as the ratio between annual dividend payment and the price of the security by the end of each month. We winsorize dividend yield at the 1% and 99% levels.

D. Estimation methodology

This Appendix provides details about the two-pass regression method used to estimate time-varying risk premia.

1. Regression framework

The first-pass time-series regression is:

$$r_{i,t} = b_{1,i}' x_{1,i,t} + b_{2,i}' x_{2,i,t} + \epsilon_{i,t} \tag{D.1}$$

$$x_{1,i,t} = (vech(X_t)', Z_{c,t-1}' \otimes Z_{i,t-1}')' \in \mathbb{R}^{d_1 = p(p+1)/2+pq}$$

$$x_{2,i,t} = [(f_{c,t}' \otimes Z_{c,t-1}'), (f_{c,t}' \otimes Z_{i,t-1}')] \in \mathbb{R}^{d_2 = K(p+q)} \tag{D.2}$$

where K is the number of factors, p is the number of common instruments, q is the number of stock-specific instruments. $X_t \in \mathbb{R}^{p \times p}$ is a symmetric matrix such that $X_{t,k,l} = Z_{c,t-1,k}^2$ if $k = l$, and $X_{t,k,l} = 2Z_{c,t-1,k}Z_{c,t-1,l}$ otherwise. $x_{1,i,t}$ contains $d_1 = p(p+1)/2 + pq$ terms, among which $vech(X_t)$ contains $p(p+1)/2$ interaction terms among common instruments and $Z_{c,t-1}' \otimes Z_{i,t-1}'$ contains pq interaction terms between common instruments and stock-specific instruments. $x_{2,i,t}$ contains $K(p+q)$ factors scaled by common and stock-specific instruments. In our specific setting, the total number of regressors in the first-pass time-series regression is $d = d_1 + d_2 = 9 + 12 = 21$.

In the second-pass cross-sectional regression, we regress $\hat{b}_1$ on $\hat{b}_3$, which is a transformation of $\hat{b}_2$:

$$\hat{b}_{1,i} = \hat{b}_{3,i}\nu_c \tag{D.3}$$

$$\hat{b}_{3,i} = \Big((N_p[B_i' \otimes I_p])', [W_{p,q}(C_i' \otimes I_p)]'\Big)' \tag{D.4}$$

$$\nu_c = vec[\Lambda_c' - F_c'] \tag{D.5}$$

where $N_p = \frac{1}{2}D_p^+(W_{p,p} + I_{p^2})$, $W_{p,q}$ is the commutation matrix such that $vec[A'] = W_{p,q}vec[A]$ and D_p^+ is the $p(p+1)/2$-by-p^2 matrix such that $vech[A] = D_p^+ vec[A]$.

We run our estimation for individual stocks in each country using global and local factors, n_c denotes number of stocks in a country and T_c denotes the total number of periods when data in country c is available.

2. The instrument selection procedure

Because there is a large number of regressors arising from interaction terms between factor loadings and factor risk-premia, in practice, we need to impose some structure and set some elements in B_i and C_i to zero. We use an instrument selection algorithm similar to Chaieb, Langlois, and Scaillet (2021).

Let $\mathbb{I}_{Bi}$ and $\mathbb{I}_{Ci}$ be $K \times p$ and $K \times q$ indicator matrices whose elements are equal to one if the corresponding elements in B_i and C_i are non-zero. $\tilde{\mathbb{I}}_{B_i}$ is the p-vector whose j^{th} element is equal to one if at least one element in the j^{th} column of B_i is not zero, $\tilde{\mathbb{I}}_{C_i}$ is the q-vector whose j^{th} element is equal to one if at least one element in the j^{th} column of C_i is not zero. $\tilde{p}_i = \tilde{\mathbb{I}}_{B_i}'\iota_p$ is the number of columns in B_i with at least one nonzero element. $\tilde{B}_i$, $\tilde{C}_i$ are obtained by removing rows in $diag(vec[\mathbb{I}_{B_i}'])$, $diag(vec[\mathbb{I}_{C_i}'])$ for which all columns are zero. $\tilde{D}_i$ is the matrix $diag(vech[diag(\tilde{\mathbb{I}}_{B_i})]) + \iota_p\iota_p' - I_p])$ in which columns with all zeros have been removed. $\tilde{E}_i$ is the matrix $diag(\iota_p \otimes \tilde{\mathbb{I}}_{C_i})$ in which columns with all zeros have been removed.

Under the restrictions on B_i and C_i, the first-pass regression is performed as:

$$r_{i,t} = b'_{1,i} x_{1,i,t} + b'_{2,i} x_{2,i,t} + \epsilon_{i,t} \tag{D.6}$$

$$x_{1,i,t} = (vech(X_t)'\tilde{D}_i, (Z'_{c,t-1} \otimes Z'_{i,t-1})\tilde{E}_i)'$$

$$x_{2,i,t} = [(f'_{c,t} \otimes Z'_{c,t-1})\tilde{B}'_i, (f'_{c,t} \otimes Z'_{i,t-1})\tilde{C}'_i]' \tag{D.7}$$

The coefficients $b_{1,i}$ and $b_{2,i}$ are transformation of the coefficients B_i, C_i, Λ_c and F_c of our linear specifications (14), (15), (16):

$$b_{1,i} = \left((\tilde{D}'_i N_p [(\Lambda_c - F_c)' \otimes I_p] \tilde{B}'_i \tilde{B}_i vec[B'_i])', ([(\Lambda_c - F_c)' \otimes I_q] vec[C'_i])' \right)' \tag{D.8}$$

$$b_{2,i} = \left((\tilde{B}_{i,c} vec[B'_i])', (\tilde{C}_{i,c} vec[C'_i])' \right)' \tag{D.9}$$

We choose the restrictions $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ using an iterative procedure:

1. Run the regression setting all elements in $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ to one.
2. Calculate the condition number of the matrix $\hat{Q}_{x,i} = \frac{1}{T_i} \sum_t I_{i,t} x_{i,t} x'_{i,t}$:

$$CN(\hat{Q}_{x,i}) = \sqrt{eig_{max}(\hat{Q}_{x,i})/eig_{min}(\hat{Q}_{x,i})} \tag{D.10}$$

3. If the condition number is above 15 or the determinant of Q_x is less than machine precision, then find the pair of regressors in $x_{i,2}$ that has the largest cross-correlation in absolute value. Of these two regressors, remove the one with the lowest absolute correlation with $r_{i,t}$ and set its corresponding element in $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ to 0.
4. We check the following condition on $\mathbb{I}_{B_i}$: the first column of B_i, i.e. constant are all selected [32]. Otherwise, we keep the regressor and look for the next regressor pair with

[32] We drop the requirement in Chaieb, Langlois, and Scaillet (2021) that each stock-specific instrument is

the highest correlation.

5. We construct new regressors using updated indicator matrices $\mathbb{I}_{B_i}$ and $\mathbb{I}_{C_i}$ and rerun the regression.

3. WLS regression and inference

The second-pass cross-sectional regression under instrument selection is:

$$\hat{b}_{1,i} = \hat{b}_{3,i}\nu \tag{D.11}$$

$$b_{3,i} = \left(\left(\tilde{D}_i' N_p[B_i' \otimes I_p]\right)', \left[W_{p,q}(C_i' \otimes I_p)\right]'\right)' \tag{D.12}$$

$$\nu_c = vec[\Lambda_c' - F_c'] \tag{D.13}$$

We use a weighted lease squares (WLS) estimator for ν_c:

$$\hat{\nu}_c^{WLS} = \hat{Q}_{b_3}^{-1} \frac{1}{n_c} \sum_i \hat{b}_{3,i}' \hat{w}_i \hat{b}_{1,i} \tag{D.14}$$

where $\hat{Q}_{b_3} = \frac{1}{n_c} \sum_i \hat{b}_{3,i}' \hat{w}_i \hat{b}_{3,i}$ and $\hat{w}_i = \mathbf{1}_i (diag[\hat{v}_i])^{-1}$ are the weights. The weights are set to be the inverse of the asymptotic variances of the standardized errors $\sqrt{T_c}(\hat{b}_{1,i} - \hat{b}_{3,1}\nu_c)$ in the cross-sectional regression for large T from a first step OLS regression. With the OLS estimator $\hat{\nu}$ we calculate standard errors $\hat{v}_i$ using the following relations: $\hat{v}_i = \tau_{i,c} C_{\hat{\nu}_c,i}' \hat{Q}_{x,i}^{-1} S_{ii} \hat{Q}_{x,i}^{-1} C_{\hat{\nu}_c,i}$. Where $\tau_i = \frac{T_c}{T_i}$, $\hat{Q}_{x,i} = \frac{1}{T_i}\sum_t I_{i,t} x_{i,t} x_{i,t}'$, $\hat{S}_{ii} = \frac{1}{T_i}\sum_t I_{i,t} \hat{\epsilon}_{i,t}^2 x_{i,t} x_{i,t}'$, $\hat{\epsilon}_{i,t} = r_{i,t} - \hat{b}_i' x_{i,t}$ and $C_{\hat{\nu}_c,i} = (E_{1,i}' - (I_{d_{1,i}} \otimes \hat{\nu}_c') J_{a,i} E_{2,i}')'$, $E_{1,i} = (I_{d_{1,i}}, \mathbf{0}_{d_{1,i},d_{2,i}})'$, $E_{2,i} = (\mathbf{0}_{d_{2,i},d_{1,i}}, I_{d_{2,i}})'$.

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selected to predict at least one factor loading, because we would like to keep more stocks in our sample to preserve meaningful dispersion in institutional ownership.

We obtain $b_{3,i}$ by invoking the following identity:

$$vec(b'_{3,i}) = J_{a,i} b_{2,i} \tag{D.15}$$

$$J_{a,i} = \begin{bmatrix} J_{1,i} & O \\ O & J_{2,i} \end{bmatrix} \tag{D.16}$$

$$J_{1,i} = W_{p(p-1)/2+\tilde{p}_i,Kp}(I_{pK} \otimes (\tilde{D}'_i N_p) \times \{I_K \otimes [(W_p \otimes I_p)(I_p \otimes vec[I_p])]\}\tilde{B}'_{i,c} \tag{D.17}$$

$$J_{2,i} = W_{pq,pK}(I_K \otimes [(I_p \otimes W_{p,q})(W_{p,q} \otimes I_p)(I_q \otimes vec(I_p))])\tilde{C}'_{i,c} \tag{D.18}$$

where $W_{p,q}$ is the $p \times q$ commutation matrix and $N_p = \frac{1}{2} D_p^+ (W_{p,p} + I_{p^2})$ where D_p^+ is the $p(p+1)/2 \times p^2$ matrix such that $vech(A) = D_p^+ vec(A)$.

The distribution of the estimator $\hat{\nu}_c^{WLS}$ is:

$$\sqrt{n_c T_c}(\hat{\nu}_c^{WLS} - \frac{1}{T_c}\hat{B}_{\nu_c} - \nu_c) \Rightarrow N(0, \Sigma_{\nu_c}) \tag{D.19}$$

Where B_ν is the bias correction:

$$\hat{B}_{\nu_c} = \hat{Q}_{b_3}^{-1} J_b \frac{1}{n_c} \sum_i \tau_i vec[E'_{2,i} \hat{Q}_{x,i}^{-1} \hat{S}_{ii} \hat{Q}_{x,i}^{-1} C_{\hat{\nu}_c} \hat{w}_i] \tag{D.20}$$

And the covariance matrix of the estimated risk premium ν_c is:

$$\hat{\Sigma}_{\nu_c} = \hat{Q}_{b_3}^{-1}\hat{S}\hat{Q}_{b_3}^{-1} \tag{D.21}$$
$$\hat{S} = \frac{1}{n_c}\sum_{i,j}\frac{\tau_i\tau_j}{\tau_{i,j}}b'_{3,i}w_iC'_{\hat{\nu}_c,i}\hat{Q}_{x,i}^{-1}\tilde{S}\hat{Q}_{x,j}^{-1}C_{\hat{\nu}_c,j}w'_{j,c}b_{3,j},\ \tau_{i,j} = \frac{T_c}{T_{ij}}$$
$$J_b = (vec[I_{d_{1,i}}]' \otimes I_{Kp})(I_{d_{1,i}} \otimes J_{a,i})$$
$$C_{\hat{\nu}_c,i} = (E'_{1,i} - (I_{d_{1,i}} \otimes \hat{\nu}'_c)J_{a,i}E'_{2,i})'$$

The bias correction $\hat{B}_{\nu_c}$ accounts for the error-in-variable problem in the first-pass regression. We use a hard thresholded estimator $\tilde{S}_{ij} = \hat{S}_{ij}\mathbf{1}_{||\hat{S}_{ij}|| \geq \kappa_{n_c,T_c}}$, where $\hat{S}_{ij} = \frac{1}{T_{ij}}\sum_t I_{i,t}I_{j,t}\hat{\epsilon}_{i,t}\hat{\epsilon}_{j,t}x_{i,t}x'_{j,t}$, $||\hat{S}_{ij}||$ is the Frobenius norm, $\kappa_{n_c,T_c} = M\sqrt{\frac{log(n_c)}{T_c}}$ is a data-dependent threshold and M is a positive number set by cross-validation.

To obtain estimates of time-varying risk premia $\hat{\Lambda}_c$, we run SUR of $f_{c,t}$ on lagged common instruments $Z_{c,t-1}$.[33]:

$$f_{c,t} = F_cZ_{c,t-1} + u_t \tag{D.22}$$

The estimator for F_c is:

$$\hat{F}_c = (\sum f_{c,t}Z'_{c,t-1})(\sum_t Z_{c,t-1}Z'_{c,t-1})^{-1} \tag{D.23}$$

Then $\hat{\Lambda}$ follows from the relation $\nu_c = vec(\Lambda'_c - F'_c)$.

[33] We impose the restriction on F_c that the loading of the world factor on country-specific instruments be zero and that the loadings of local factors on the world instrument be zero.

The asymptotic distribution of Λ_c is: $\sqrt{T_c} vec[\hat{\Lambda}'_c - \Lambda'_c] \Rightarrow N(0, \Sigma_{\Lambda_c})$, where

$$\Sigma_{\Lambda_c} = (\mathbb{I}_K \otimes Q_z^{-1})\Sigma_u(\mathbb{I}_K \otimes Q_z^{-1}) \tag{D.24}$$

$$\Sigma_u = E[u_t u'_t \otimes Z_{c,t-1} Z'_{c,t-1}]$$

$$Q_z = E[Z_{c,t-1} Z'_{c,t-1}]$$

The asymptotic distribution of Λ_c is dominated by the asymptotic distribution of F_c because $\hat{\nu}_c$ has a faster convergence rate than F_c, as is explained in Gagliardini, Ossola, and Scaillet (2016).

Finally, the time-varying risk premia is calculated as $\hat{\lambda}_{c,t} = \hat{\Lambda}_c Z_{c,t-1}$. The time-varying risk premia of each individual stock is calculated as $\mu_{i,t} = \hat{\beta}'_{i,t} \mu_{c,t}$. Where $\hat{\beta}_{i,t}$ is obtained from first-pass estimation $\hat{b}_i$ though (D.9) and (14).

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Table I. Summary statistics

This table reports for each country the total number of stocks (n), the average proportion of institutional stocks to the country's aggregate market capitalization (proportion), annualized average returns and volatilities of the institutional local factor (inst local) and the retail local factor (retail local). For each month, institutional stocks are stocks whose global institutional ownership is above the 50th percentile of all stocks in its country and higher than 1%. The institutional local factor is constructed as the residual from a 36-month rolling window regression of the attainable domestic market portfolio onto the foreign institutional portfolio. The retail local factor is constructed as the residual from a 36-month rolling window regression of the retail portfolio onto the domestic institutional portfolio. The sample period is monthly from January 2000 to December 2020.

			Annualized average return (%)		Annualized volatility (%)	
			A: Developed markets			
Country	n	proportion	inst local	retail local	inst local	retail local
Australia	3209	0.854	3.239	-2.776	11.888	7.713
Austria	183	0.800	0.108	2.701	12.595	9.876
Belgium	290	0.812	-0.048	2.052	11.990	11.798
Canada	3622	0.926	1.562	-0.423	10.237	11.074
Denmark	378	0.816	7.052	-2.629	12.340	11.880
Finland	256	0.812	-3.014	4.769	16.718	18.504
France	1556	0.875	-2.048	2.223	9.006	8.509
Germany	1352	0.857	-4.351	4.563	9.256	10.886
Hong Kong	1924	0.869	4.099	-11.752	13.384	15.535
Ireland	176	0.820	2.245	-7.423	11.337	38.427
Israel	890	0.750	2.068	-0.026	15.582	15.135
Italy	679	0.881	-3.860	-8.395	13.150	11.868
Japan	4842	0.881	-1.450	-2.433	12.903	7.772
Netherlands	396	0.727	0.580	-5.190	10.602	11.700
New Zealand	273	0.741	6.739	-0.820	12.693	10.800
Norway	530	0.775	2.850	2.599	14.186	13.507
Portugal	114	0.883	-4.885	-3.554	13.806	21.639
Singapore	1006	0.834	2.433	-2.620	13.582	9.534
Spain	382	0.814	-0.458	-5.527	13.897	12.387
Sweden	995	0.879	-0.945	6.631	10.733	10.392
Switzerland	478	0.881	2.496	-3.762	9.234	11.509
UK	4333	0.843	-2.726	-2.482	7.549	7.108
US	15630	0.911	0.736	-5.019	7.530	13.921
			A: Emerging markets			
country	n	proportion	inst local	retail local	inst local	retail local
Brazil	526	0.657	6.916	-8.304	30.898	22.749
Chile	270	0.484	-5.052	7.072	20.929	12.733
China	5566	0.334	13.607	-5.355	17.531	22.553

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			Annualized average return (%)		Annualized volatility (%)	
Greece	464	0.719	-12.358	-6.588	26.010	16.203
India	4920	0.764	6.065	-1.609	23.338	13.591
Indonesia	805	0.649	12.245	-7.728	29.545	17.407
South Korea	3123	0.822	5.041	-12.964	18.253	13.053
Malaysia	1429	0.756	4.314	-2.040	16.005	8.925
Mexico	217	0.750	3.274	-0.292	15.271	11.611
Philippines	320	0.716	8.937	-1.341	19.408	13.528
Poland	1110	0.810	-0.364	-0.259	20.180	12.378
South Africa	831	0.782	6.393	-2.247	19.649	11.692
Taiwan	2494	0.802	2.055	-5.323	18.263	10.909
Thailand	956	0.677	12.596	-7.085	21.589	13.258
Turkey	524	0.777	1.634	-3.238	33.559	21.245

Table II. Price of covariance risk estimated from unconditional Fama-MacBeth regression of individual stocks returns.

This table reports our unconditional estimate of the price of covariance risk of the attainable world market factor, the institutional local factor and the retail local factor. We report the estimated price of covariance risk as the average across cross-sections. We report Newey and West (1987) standard errors that account for heteroskedasticity and serial correlation.

$$E[r_i] = \alpha + \lambda^{\hat{W}} cov(r_i, r_{\hat{W}}) + \lambda^{ilocal} cov(r_i, f^{ilocal}) + \lambda^{rlocal} cov(r_i, f^{rlocal})$$

In each month, we use a 36-month rolling regression to estimate each covariance. Then for each cross-section, we regress stock returns on the covariance estimates. For each country, we also report underneath country name the average number of stocks in each cross-section. We also report the average OLS adjusted R^2 in the first row and the GLS R^2 in the second row as measures of model fit. * indicates that the coefficient is significant at the 10% level, ** indicates that the coefficient is significant at the 5% level and *** indicates that the coefficient is signficant at the 1% level.

A: Developed markets					
Country	α	$\lambda^{\hat{W}}$	λ^{ilocal}	λ^{rlocal}	R2OLS/GLS
Australia	0.002	0.459	1.514***	0.495	0.030
1734	(0.588)	(0.782)	(3.555)	(0.945)	0.027
Austria	-0.000	-1.388	8.163***	5.778***	0.266
107	(-0.127)	(-0.906)	(4.387)	(4.813)	0.264
Belgium	0.002	0.611	3.360***	-2.470*	0.209
174	(0.523)	(0.419)	(3.087)	(-1.829)	0.202
Canada	-0.006***	1.232**	2.330***	0.881**	0.036
1796	(-3.063)	(2.383)	(3.176)	(2.014)	0.034
Denmark	0.005	-1.099	0.142	2.181**	0.108
209	(1.501)	(-1.103)	(0.128)	(2.316)	0.111
Finland	0.002	0.976	4.028***	0.885	0.120

148	(0.429)	(0.683)	(2.953)	(1.214)	0.126
France	0.004	-0.974	1.706***	0.082	0.054
881	(1.205)	(-1.247)	(2.773)	(0.098)	0.048
Germany	-0.005	-0.951	4.270***	0.981	0.063
822	(-1.102)	(-1.165)	(4.784)	(1.600)	0.060
Hong Kong	-0.001	-0.534*	1.201***	0.224	0.036
1165	(-0.188)	(-1.716)	(3.371)	(0.476)	0.034
Ireland	0.002	1.984*	6.422***	-1.101*	0.183
98	(0.724)	(1.837)	(3.036)	(-1.965)	0.199
Israel	-0.001	0.219	1.107**	2.276***	0.052
474	(-0.386)	(0.536)	(2.063)	(7.092)	0.049
Italy	0.003	-0.578	1.347	-2.181*	0.090
329	(1.600)	(-0.886)	(1.093)	(-1.940)	0.088
Japan	0.002	-0.717	2.089***	1.216	0.066
3515	(1.092)	(-1.308)	(4.996)	(0.823)	0.057
Netherlands	0.003**	-0.411	-0.169	-1.372	0.109
220	(2.448)	(-0.314)	(-0.150)	(-0.943)	0.111
New Zealand	0.004	-0.999	1.703**	-1.173	0.119
147	(0.730)	(-0.898)	(2.543)	(-1.167)	0.124
Norway	0.004	-1.126	-0.549	2.132***	0.104
231	(1.185)	(-1.630)	(-0.607)	(2.681)	0.104
Portugal	0.005	-1.511	-1.277	2.352***	0.194
63	(1.504)	(-1.204)	(-1.448)	(3.239)	0.220
Singapore	0.005***	-0.002	0.001	1.362**	0.072
608	(2.991)	(-0.002)	(0.002)	(2.285)	0.068
Spain	0.004	-1.422**	2.326***	0.366	0.097
189	(1.096)	(-2.120)	(2.748)	(0.244)	0.103
Sweden	-0.001	1.158**	3.424***	1.445	0.065
444	(-0.288)	(2.481)	(4.332)	(1.629)	0.062
Switzerland	0.001	0.236	6.335***	-0.256	0.114
299	(0.658)	(0.187)	(4.803)	(-0.268)	0.109
UK	-0.003	0.663	0.865	-2.689***	0.025
2251	(-1.129)	(1.615)	(1.642)	(-3.679)	0.023
US	-0.020***	3.120***	-1.000	-3.188***	0.026
8738	(-5.953)	(3.049)	(-0.943)	(-3.705)	0.022
			Emerging markets		
Country	α	$\lambda^{\hat{W}}$	λ^{ilocal}	λ^{rlocal}	R2OLS/GLS
Brazil	0.011	-0.471	0.700	1.846***	0.131
271	(1.536)	(-0.488)	(1.132)	(2.853)	0.104

Chile	0.009*	0.847	-0.771	1.983**	0.070
180	(1.833)	(1.199)	(-1.387)	(2.039)	0.056
China	-0.003	-0.143	2.103**	1.103*	0.128
2625	(-1.045)	(-0.344)	(2.036)	(1.895)	0.105
Greece	-0.004	2.339**	-0.507	-0.114	0.097
298	(-1.467)	(2.142)	(-1.227)	(-0.140)	0.096
India	0.007*	-0.462	1.515***	0.180	0.050
2321	(1.738)	(-1.132)	(5.463)	(0.376)	0.043
Indonesia	0.009**	-0.661	0.857***	2.189**	0.059
431	(2.335)	(-1.279)	(3.662)	(2.563)	0.055
Malaysia	0.004**	-0.460	0.788***	-0.393	0.043
977	(2.514)	(-1.030)	(2.663)	(-0.553)	0.041
Mexico	0.007*	-1.334	1.808**	3.103	0.128
126	(1.707)	(-0.901)	(2.100)	(1.312)	0.131
Philippines	0.006	0.406	0.803**	1.781***	0.077
231	(0.962)	(0.817)	(2.174)	(2.746)	0.079
Poland	-0.001	-0.163	0.476	2.152***	0.067
507	(-0.557)	(-0.231)	(1.397)	(3.807)	0.063
South Africa	0.000	0.859**	0.585	-2.004*	0.053
420	(0.048)	(2.314)	(1.028)	(-1.927)	0.051
South Korea	-0.008**	1.469***	1.548***	0.726	0.083
1726	(-2.368)	(2.764)	(4.641)	(0.710)	0.068
Taiwan	-0.001	1.410***	0.642**	0.727	0.056
1463	(-0.410)	(3.491)	(2.400)	(1.175)	0.049
Thailand	0.009***	-0.415	0.072	-0.784	0.081
570	(3.751)	(-0.792)	(0.168)	(-0.873)	0.061
Turkey	0.005	1.384***	0.654***	-0.006	0.082
345	(1.332)	(2.690)	(3.304)	(-0.020)	0.080

Table III. Attainable world market, institutional local and retail local risk premia estimated from the Gagliardini, Ossola, and Scaillet (2016) conditional two-pass regression.

This table presents for each country our estimates of the loadings of time-varying risk premia on common instruments Λ_c in the linear specification of time-varying risk premia in (15) whose components is given in (19). We also report the number of stocks (n) used in the second-step cross-sectional regression. We use a constant, world dividend yield DY_t and country dividend yield $DY_{c,t}$ as our common instruments. As is explained in Section III.B, we impose the restriction that $\Lambda_{DY}^{ilocal} = \Lambda_{DY}^{rlocal} = 0$ in our estimation. Because we also impose the restriction that $\Lambda_{DY}^{\hat{W}} = F_{DY}^{\hat{W}}$ and the conditional mean of the global factor does not load on local dividend yield, $\Lambda_{DY}^{\hat{W}}$ is the same across countries. In this table, we only report the coefficients of the constant Λ_0 and of the local dividend yield Λ_{DY_c}. We first obtain estimates for the risk-premium ν_c from the second-pass regression, then we estimate F_c using a SUR of $f_{c,t}$ on $Z_{c,t-1}$. We then obtain Λ_c through the relation $\nu_c = vec(\Lambda_c' - F_c')$. The covariance matrix of Λ_c is given in (D.24). * indicates that the coefficient is significant at the 10% level, ** indicates that the coefficient is significant at the 5% level and *** indicates that the coefficient is signficant at the 1% level. All numbers are reported in annualized percentage terms.

A: Developed markets

		Attainable world		Institutional local		Retail local	
Country	n	Λ_0	Λ_{DY_c}	Λ_0	Λ_{DY_c}	Λ_0	Λ_{DY_c}
Australia	1681	3.041	3.386 **	7.395	−4.665	3.622	−1.418
Austria	75	4.620 *	−3.278 *	4.273	−5.671 *	7.967 **	−3.048
Belgium	118	−2.172	−9.507 ***	24.925 ***	16.054 ***	−9.511 *	5.271 **
Canada	1529	8.627***	−3.071 *	−1.686	−5.901	−2.510	−0.032
Denmark	158	3.068	1.444	12.044 ***	−4.986	−3.438	4.922 *
Finland	122	11.120***	−0.787	−7.710	5.000	0.672	−1.906
France	717	−0.883	−0.190	6.384	−1.959	6.443	−0.013
Germany	721	6.005 **	−2.636	−1.205	3.753	2.490	1.068
Hong Kong	1141	2.571	0.815	6.429 *	4.386	7.837 *	2.006
Ireland	77	12.974***	0.424	−7.984 **	−2.718	20.742***	−2.501
Israel	453	9.429***	−0.691	−2.760	−0.403	7.550 **	4.008
Italy	282	1.013	−0.299	1.826	−0.629	−4.378	−4.487
Japan	2670	3.919	−3.483 **	1.051	2.638	5.406	0.584
Netherlands	185	4.519 *	−0.934	3.509	−1.943	−5.113	3.985
New Zealand	120	11.733***	3.978 **	0.109	−1.387	2.111	−0.680
Norway	202	7.317 **	−5.357 **	4.560	−3.770	1.602	3.178
Portugal	44	8.690***	−15.832 ***	−7.387	17.284 ***	−4.041	−13.193 **
Singapore	562	6.882 **	−0.637	−1.695	0.577	7.330	−1.096
Spain	155	5.071 *	−3.796 *	1.957	−3.739	−9.298	−8.968 **
Sweden	393	4.188	2.892	6.175	0.449	5.753	−7.703 **

Switzerland	222	2.804	−3.326 *	11.680 *	−1.844	−4.501	1.341
UK	1918	6.134 **	−3.255 **	−1.018	−1.858	2.285	0.906
US	7739	5.996 **	−7.488 ***	2.646	5.050	0.261	0.534
Average		5.507		2.762		1.708	

B: Emerging markets

		Attainable world		Institutional local		Retail local	
Country	n	Λ_0	Λ_{DY_c}	Λ_0	Λ_{DY_c}	Λ_0	Λ_{DY_c}
Brazil	245	8.432***	4.762	4.062	−1.821	1.700	−4.555
Chile	131	7.508 **	−1.025	2.336	−4.646	3.333	4.573 *
China	2868	8.056***	3.358	6.055	−1.572	5.375 *	−7.377 *
Greece	292	6.674 **	−5.010 *	−17.841 ***	12.269 ***	11.812***	4.837
India	2322	1.068	−0.747	16.017 ***	16.266 ***	−1.505	5.525 **
Indonesia	420	−1.955	4.834	28.115 ***	5.117	−8.007 **	1.958
South Korea	1827	7.015 **	2.067	3.444	5.478 **	5.045 *	−0.405
Malaysia	937	0.366	1.628	7.235 *	4.495 *	−0.801	−0.804
Mexico	104	14.508***	−3.463 *	−7.007 *	−3.945	4.109	5.294 **
Philippines	209	4.967 *	11.228 ***	8.638 **	1.471	6.674	2.313
Poland	610	5.114 *	4.543 **	−4.001	−6.155	7.724 *	−2.328
South Africa	323	−2.345	−5.644 **	17.102 ***	8.657 ***	−0.976	−5.625 **
Taiwan	1508	0.302	8.542 ***	5.815	0.918	3.437	−0.207
Thailand	512	−0.103	10.121 ***	19.581 ***	−12.402 ***	−6.134	3.119
Turkey	374	3.789	6.771 *	4.549	−6.626	7.888 **	−5.909
Average		4.226		6.273		2.645	

Table IV. Average model-implied risk premia across firms by country

This table reports average model-implied risk premia across individual stocks by country. For stock i from country c in month t, the total model-implied risk premium is calculated as $\mu_{i,t} = \beta'_{i,t}\mu_{c,t}$, the attainable world market risk premium is calculated as $\mu^{\hat{W}}_{i,t} = \beta^{\hat{W}}_{i,t}\mu^{\hat{W}}_{c,t}$, the institutional local risk premium is calculated as $\mu^{ilocal}_{i,t} = \beta^{ilocal}_{i,t}\mu^{ilocal}_{c,t}$ and the retail local risk premium is calculated as $\mu^{rlocal}_{i,t} = \beta^{rlocal}_{i,t}\mu^{rlocal}_{c,t}$. For each country at each time, we calculate the equal-weighted risk premia across individual stocks and report the time-series average. All numbers are reported in annualized percentage terms.

A: Developed markets				
Australia	5.211	6.949	4.238	16.397
Austria	4.818	2.656	3.471	10.944
Belgium	-1.732	12.172	-3.530	6.910
Canada	10.428	-2.341	-1.051	7.036
Denmark	3.231	7.665	-0.969	9.928
Finland	11.876	-3.530	0.290	8.636
France	-0.551	4.234	3.406	7.090
Germany	6.233	-0.428	1.143	6.949
Hong Kong	3.460	6.050	5.718	15.228
Ireland	14.498	-3.675	1.994	12.816
Israel	8.126	-2.017	5.311	11.420
Italy	1.620	1.574	-1.414	1.780
Japan	2.893	0.871	5.129	8.893
Netherlands	5.143	1.925	-1.232	5.835
New Zealand	12.753	-0.026	0.871	13.598
Norway	9.795	3.009	0.442	13.246
Portugal	7.439	-6.688	0.402	1.153
Singapore	8.411	-1.563	6.490	13.338
Spain	5.237	1.427	-3.361	3.303
Sweden	5.279	4.595	1.047	10.922
Switzerland	2.745	8.185	-1.395	9.535

UK	5.925	-1.154	1.007	5.778
US	4.885	1.438	-0.006	6.317
Average	5.278	1.426	1.818	8.522

B: Emerging markets

Country	Attainable world	Istitutional local	Retail local	Total risk premium
Brazil	12.364	3.231	1.312	16.907
Chile	6.326	1.748	2.871	10.945
China	7.833	3.869	4.791	16.492
Greece	7.975	-11.910	9.927	5.993
India	2.150	15.107	-1.542	15.716
Indonesia	-1.987	21.179	-4.782	14.410
South Korea	9.744	3.345	4.587	17.676
Malaysia	0.587	7.520	-0.940	7.168
Mexico	15.613	-5.506	1.549	11.656
Philippines	4.620	7.147	2.253	14.020
Poland	7.426	-3.198	5.454	9.682
South Africa	-2.652	12.866	-0.391	9.823
Taiwan	0.609	5.598	3.221	9.429
Thailand	-0.121	15.682	-5.293	10.268
Turkey	3.722	3.975	3.865	11.562
Average	4.519	6.636	1.968	13.122

Table V. How institutional ownership affects global and local risk premia in developed markets.

This table presents regression of firm-level model-implied total, attainable world market, institutional local and retail local premia on institutional ownership (IO) and firm level controls. For each regression, we consider two alternative specifications, one including country-level institutional ownership ($CountryIO$) and the intermediary capital ratio of He, Kelly, and Manela (2017) (CR), the other including country-time fixed effects. All standard errors are clustered at the firm level.

$$y_{i,t} = \beta_1 IO_{i,t-1} + \beta_2 \rho_i + \beta_3 X_{i,t-1} + \beta_4 CountryIO_{c,t-1} + \beta_5 CR_{t-1} + \alpha_{c,t} + \epsilon_{i,t},\ y \in \{\mu_{i,t}, \mu_{i,t}^{world}, \mu_{i,t}^{ilocal}, \mu_{i,t}^{rlocal}\}$$

	Total		World		Institutional local		Retail local	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
IO	0.063***	0.036***	0.046***	0.014***	0.029***	0.019***	-0.012***	0.004***
	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)	(0.001)	(0.001)
ρ	0.163***	0.177***	0.100***	0.129***	0.039***	0.038***	0.024***	0.010***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)
logmv	-0.009***	-0.007***	-0.005***	-0.003***	-0.001***	-0.001***	-0.002***	-0.002***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
bm	-0.002***	-0.002***	0.000	-0.001***	-0.002***	-0.000	0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
dy	-0.295***	-0.360***	-0.224***	-0.263***	-0.046***	-0.043***	-0.025***	-0.053***
	(0.012)	(0.011)	(0.010)	(0.007)	(0.008)	(0.006)	(0.005)	(0.005)
CountryIO	-0.219***		-0.044***		-0.034***		-0.141***	
	(0.005)		(0.004)		(0.004)		(0.002)	
CR	-2.576***		-1.718***		-0.481***		-0.377***	
	(0.021)		(0.018)		(0.013)		(0.008)	
Observations	3,220,189	3,220,189	3,220,189	3,220,189	3,220,189	3,220,189	3,220,189	3,220,189
R-squared	0.185	0.412	0.173	0.640	0.019	0.296	0.088	0.359
Country-time FE	N	Y	N	Y	N	Y	N	Y

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table VI. How institutional ownership affects global and local risk premia in emerging markets.

This table presents regression of firm-level model-predicted total, attainable world market, institutional local and retail local premia on institutional ownership (IO) and firm level controls. For each regression, we consider two alternative specifications, one including country-level institutional ownership ($CountryIO$) and the intermediary capital ratio of He, Kelly, and Manela (2017) (CR), the other including country-time fixed effects. All standard errors are clustered at the firm level.

$$y_{i,t} = \beta_1 IO_{i,t-1} + \beta_2 \rho_i + \beta_3 X_{i,t-1} + \beta_4 CountryIO_{c,t-1} + \beta_5 CR_{t-1} + \alpha_{c,t} + \epsilon_{i,t},\ y \in \{\mu_{i,t}, \mu_{i,t}^{world}, \mu_{i,t}^{ilocal}, \mu_{i,t}^{rlocal}\}$$

	Total		World		Institutional local		Retail local	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
IO	-0.128***	-0.081***	0.003	0.063***	-0.003	-0.043***	-0.128***	-0.102***
	(0.011)	(0.009)	(0.010)	(0.007)	(0.010)	(0.005)	(0.006)	(0.004)
ρ	0.093***	0.217***	0.039***	0.100***	0.107***	0.125***	-0.053***	-0.008***
	(0.005)	(0.004)	(0.003)	(0.003)	(0.005)	(0.004)	(0.003)	(0.002)
logmv	0.007***	-0.003***	0.007***	-0.002***	-0.007***	-0.002***	0.006***	0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
bm	0.014***	0.000	0.009***	0.002***	0.003***	-0.000	0.001***	-0.001***
	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
dy	-0.237***	-0.347***	-0.142***	-0.086***	-0.003	-0.194***	-0.092***	-0.066***
	(0.017)	(0.014)	(0.013)	(0.009)	(0.016)	(0.010)	(0.009)	(0.006)
CountryIO	0.211***		0.029***		-0.099***		0.281***	
	(0.014)		(0.011)		(0.013)		(0.008)	
CR	-3.466***		-3.869***		0.087***		0.317***	
	(0.028)		(0.025)		(0.018)		(0.009)	
Observations	1,790,675	1,790,675	1,790,675	1,790,675	1,790,675	1,790,675	1,790,675	1,790,675
R-squared	0.153	0.651	0.273	0.752	0.033	0.655	0.045	0.626
Country-time FE	N	N	N	Y	N	Y	N	Y

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1